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Agricultural Economics Research

OCTOBER 1982 VOL. 34 NO. 4

A stylized, high-contrast illustration in white and orange. It features two large corn cobs with detailed kernels, positioned diagonally across the upper half of the cover. Below the cobs, there are abstract, angular shapes representing a field or landscape, also in white and orange. The background is a solid orange color.

**Shares or cash:
how to pay the rent?**

Agricultural Economics Research

A Journal of the U.S. Department of Agriculture • Economic Research Service

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The Secretary of Agriculture has determined that the publication of this periodical is necessary in the transaction of public business required by law of this Department. Use of funds for printing this periodical has been approved by the Director, Office of Management and Budget, through March 31, 1987.

This periodical is available for sale from the following source:

Superintendent of Documents
U.S. Government Printing Office
Washington, D.C. 20402

Microfiche copies are available from two different sources:

Infodata International, Inc.
Suite 4602
175 East Delaware Place
Chicago, Ill. 60611

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In This Issue

Sometimes it pays to look at old problems in new ways and, instead of enjoying the comfort of the apparent correspondence of accepted theory and accumulated empirical evidence, to puzzle about the anomalies that accepted theory fails to explain. "To be accepted," said Thomas Kuhn in *The Structure of Scientific Revolutions*, "a theory must seem better than its competitors, but it need not, and in fact never does, explain all of the facts with which it can be confronted."

Each of the four articles in this issue addresses problems that have been discussed before in these pages and elsewhere. Each author exposes an anomaly in an accepted paradigm. This is good research technique; one example of failure is sufficient to falsify a theory. Each author then enthusiastically proposes and defends an alternative procedure. The alternative appears to work in the instances cited. You and I might do well to recall, as we read of these new ideas, that each one need not "explain all the facts with which it can be confronted."

Baron returns to the issue of whether owner-operated and cash-rent farms are more efficient than share-rent farms. He finds that the old answers to this question may be wrong when risk is taken into account. He goes further and demonstrates that the issue cannot be resolved by pure theory. It is, instead, an empirical question. The answer depends on such variables as relative costs, relative returns, degree of risk, and intensity of risk aversion.

Belongia and Dickey demonstrate that the usual procedures used to filter out serial-correlated disturbances from time series in order to test for causality will occasionally fail. The authors propose an alternative filter that uses an autoregressive-moving-average model, and they illustrate with a successful application.

Lutton addresses the problem of statistical estimates of production functions and of derived input demand curves. He explains the limitations of traditional, direct procedures for estimating these equations and explains how developments in the theory of duality facilitate obtaining them. Because duality estimates rely on a different data set than do direct estimates, duality theory is sometimes the only feasible procedure for circumventing data limitations.

LeBlanc is concerned with estimates of shares. In such problems, the dependent variable is constrained to lie within the unit interval. Ordinary regression may provide estimates outside the required interval. Ordinary binary choice models only recognize the endpoints of the interval (0 and 1). LeBlanc illustrates the application of a multivariate logit framework to estimate and explain input cost shares for agriculture.

Clark Edwards

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The Effects of Tenancy and Risk on Cropping Patterns: A Mathematical Programming Analysis

By Donald Baron*

Abstract

Most analyses of allocative efficiency under different forms of agricultural tenure—share tenancy, fixed cash tenancy, and owner cultivation—employ single-product models of production. These models show that risk sharing encourages share tenants to produce as much as or more than equally risk-averse owner-operators and cash tenants. However, when risk and risk aversion are introduced into multiproduct linear programming models, relative allocative efficiency under share tenancy may decline. The result depends on the relative production costs and the relative risk premiums of the different products.

Keywords

Farm tenancy, risk, allocative efficiency

The 1978 Census of Agriculture reports that nearly 40 percent of all land in farms in the United States is rented and that approximately 86 percent of this rented acreage is owned by nonoperator landlords (45).¹ Studies by Reinsel and Johnson (28), Johnson (21), and Reiss (29, 30, 31) suggest that the contributions landlords make to the management of the land they rent out has declined in recent years. An increasing percentage of tenants simultaneously rents several tracts of land from several landlords rather than a single tract from a single landlord. Contributions made by individual landlords to total rented acreage operated by tenants have declined, as have opportunities and incentives for landlords to contribute entrepreneurial skills and to help finance production costs (28, 29, 30, 31). Tenants have expressed greater willingness to operate without assistance from landlords. Many have indicated they prefer arrangements that provide them primary managerial control over farming operations so that they can coordinate their owned and rented resources effectively (6, 7, 22, 28, 29).

The increasing frequency of rental arrangements that allocate most managerial responsibilities to tenants has renewed interest in an issue that Alfred Marshall first raised: Is the efficiency of resource allocation lower on tenant-operated than on owner-operated farmland? Many economists have argued that tenants who operate with little managerial or cost-sharing assistance from nonoperator landlords, but who pay landlords fixed cash rents, will farm just as efficiently as will owner-operators. However, if these tenants pay share rents, they

will farm less efficiently than will either owners or cash tenants (1, 4, 11, 13, 14, 15, 16, 20, 23, 24, 26, 32, 33).

Most analyses of farm tenancy have employed single-input, single-product models of production. Owner-operators and fixed cash tenants receive the entire value of the marginal product from each unit of input, and they employ inputs up to an amount at which marginal value product equals marginal input cost. Share tenants pay a share of the marginal value product as rent. The level of input employment at which their own share of marginal value product equals marginal input cost is lower than the equilibrium level under owner operation and cash tenancy. The single-input, single-product model, therefore, predicts that share tenants will produce less under equilibrium than do owner-operators and cash tenants.²

Other studies of farm tenancy have applied multi-input, multiproduct, linear programming models. Within a range of output-price ratios, share tenants maximize their own net revenues, according to these models, by producing greater amounts of lower cost products and smaller amounts of higher cost products than owner-operators and cash tenants produce. Because output combinations chosen by owners and cash tenants always maximize total net revenues for the entire farm, combinations chosen by share tenants yield less than the maximum net farm revenues. The discrepancy measures the loss of efficiency attributable to the share rental arrangement within the linear programming model (10, 14, 17, 23, 25).

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¹ Italicized numbers in parentheses refer to items in the References at the end of this article.

² Many economists recommend that landlords share variable costs with their share tenants in the same proportion that they share output. As the marginal value received by the tenants will then equal the share of input cost they must pay, they will be encouraged to employ as much of the input as owners and cash tenants employ (1, 2, 3, 15, 16, 18, 23, 27).

Incorporating Risk and Risk Aversion into the Single- and Multiproduct Models

Previous multiproduct programming models of farm tenancy have either ignored production risks or have assumed that all farm operators are risk neutral. They have not considered how risk and risk aversion might affect predictions of tenancy-related differences in farm planning.

A few presentations of single-input, single-product models have addressed this question. Sutinen (34, 35) and Baron (5) showed that, although production by risk-averse operators declines as risk levels increase, this reduction is smaller under share tenancy than under owner operation and cash tenancy. Risk-sharing increases share tenants' production relative to owner-operator and cash tenants' production. The excess of owner-operator/cash tenants' production over share tenants' production that occurs under risk neutrality is, therefore, always reduced, or possibly eliminated, under risk and risk aversion (5; 19, p. 24). If the negative effect of marginal product sharing under share tenancy is overcome through landlord supervision or through contractual specification of tenants' obligations to employ variable inputs, risk sharing may encourage share tenants to produce even more than owners and cash tenants produce (35, pp. 617-19).

However, risk sharing may not have the same unequivocally positive impact on the efficiency of share tenancy relative to owner operation and to cash tenancy when it is examined within a multi-input, multiproduct programming model. To introduce risk and risk aversion into such a model, several assumptions need to be made about the attitudes of farm operators toward risk and toward changes in risk levels associated with changes in production levels. First, assume operators measure risk as a dispersion of gross revenue per acre of each crop around expected gross crop revenue per acre. Moreover, operators assign a cost, or risk premium, to this risk level. The premium equals some constant percentage of expected gross crop revenue per acre. Thus, the marginal rate of increase in the risk premium that results from an increase in production will be constant at all levels of production. Risk premiums for each crop will be included in the objective function as constant costs per acre of output, so that the function will remain linear.

Price maps for alternative output combinations can display the results of parametric programming of the objective function. Two sets of maps are presented in this study. The first is derived for a risk-neutral owner-operator and cash tenant and for a risk-neutral share tenant through maximization of the standard risk-neutral objective function presented in previous studies. This set is compared with a second set of price maps derived for risk-averse owner-operators and cash tenants, and for risk-averse share tenants through maximization of the risk-adjusted or certainty-equivalent objective function defined here. The comparison shows that the introduction of risk and

risk aversion into the programming model may either increase or decrease the likelihood that an inefficient share tenant plan will be chosen, depending on the relative prices, relative input costs, and relative risk premiums per acre of soybeans and corn. That is, contrary to the conclusion of other studies reviewed above, risk and risk aversion may either augment or offset the negative effect of product-sharing on the efficiency of share tenancy relative to owner operation and cash tenancy.

Price Mapping Under Risk Neutrality

The programming model considers alternative production plans on a 200-acre farm in Illinois that produces soybeans or corn. Cost estimates for this farm are available from the Firm Enterprise Data System (FEDS) developed by the Economic Research Service and Oklahoma State University (12).

The most recent FEDS estimates, based on data from the 1977 crop year, suggest that variable costs per acre of corn are approximately twice as high as those for soybeans. These estimates are derived from a production function which utilized approximately 30 inputs. For convenience, the programming model used here assumes that land is the only resource in limited supply.

A simplified linear programming model of farm planning by a risk-neutral operator can be represented by the equation below.

Maximize:

$$I = (rp_1 - c) a_1 + (rp_2 - hc) a_2$$

subject to:

$$a_1 + a_2 = 200 \text{ acres}$$

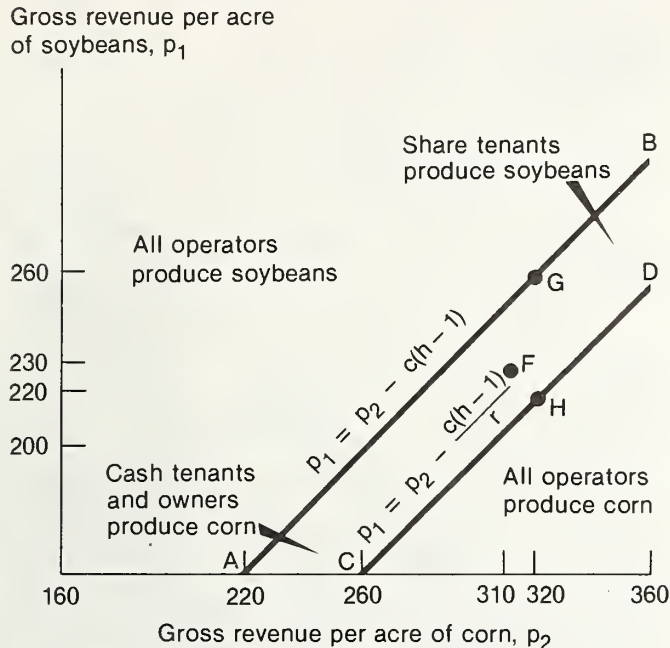
where I is net farm income; p_1 and p_2 are gross revenues per acre of soybeans and corn, respectively; r is the share rental rate, which equals 1 for owner-operators and cash tenants and some fraction less than 1 for share tenants; a_1 and a_2 are total soybean and corn acreages; c and hc are costs per acre of soybeans and corn; and h is the ratio of per acre costs of corn to soybeans. The constraint equation indicates that 200 acres of land are available.

Parametric variation of the coefficients, $rp_1 - c$ and $rp_2 - hc$, defines price maps which indicate the combinations of soybean and corn acreages that yield maximum profits for farm operators at any given ratio of gross soybean revenue to gross corn revenue. As land is the sole resource constraint, only two plans may be optimal at any given ratio: "soybeans only" or "corn only."

The price map for the risk-neutral owner-operator and cash tenant is represented by AB in figure 1. In this situation,

Figure 1

Price Map for Risk-Neutral Farmers



$r = 1$. The axes measure gross revenues per acre of corn and soybeans. For all per-acre revenue combinations which fall to the left of AB, owner-operators and cash tenants receive higher net revenues per acre from soybeans than from corn. Thus, they produce soybeans only. For all combinations to the right of AB, they receive more from corn than from soybeans, and they produce corn only.

For risk-neutral share tenants, assume that r equals 0.6 and that the tenants receive no cost sharing from landlords, so that their variable costs remain at c per acre of soybeans and at hc per acre of corn.³ The share tenants' price map is defined by line CD, which is the new boundary between the soybeans-only and corn-only farm plans. Revenue combinations within the area bounded by AB and CD define a different

³ This assumption may be challenged on grounds that proportionate cost sharing by U.S. landlords is quite common. However, Berry (6, 8, 9) and Harwell and Strickland (14) report that, in the Great Plains Region, many share-rental landlords and tenants avoid cost sharing because they consider it a partnership arrangement which demands precisely the type of landlord participation in farm management they want to avoid. These landlords and tenants typically choose share leases which assign most management responsibilities to tenants, including responsibility for financing variable costs. Reiss (29, 30, 31) reports that even in Illinois, where cost sharing has been widely practiced, a trend has developed toward greater use of net share leases—that is, share leases which assign all variable costs to tenants. Again, the motivation may be greater preference for leases which assign landlords little role in farm management or in the financing of production costs.

optimal plan for share tenants than for owner-operators and cash tenants. Combinations falling to the left of AB and to the right of CD result in the adoption of the same plan by all operators.

Maps AB and CD define ratios of gross corn to gross soybean revenue at which net soybean and corn revenues are equal. CD is further right than AB, indicating that share tenants require a higher ratio of corn revenue to soybean revenue before they will switch from soybeans to corn. To determine why CD is further right, observe that the equality between net soybean and corn revenues per acre is:

$$rp_1 - c = rp_2 - hc$$

which implies that:

$$p_1 = p_2 - c(h-1)/r \quad (1)$$

Under owner operation and cash tenancy, $r = 1$; equation (1) indicates that gross revenue per acre of corn (p_2) must exceed gross revenue per acre of soybeans (p_1) at all points on AB by exactly the amount that cost per acre of corn exceeds cost per acre of soybeans ($c(h-1)$). Under share tenancy, $r = 0.6$; equation (1) indicates that gross corn revenues per acre must exceed gross soybean revenues per acre at all points on CD by approximately $1/0.6$ (or 1.67) times the cost differential.

The price maps also indicate why allocative efficiency is lower under share tenancy than under owner operation and cash tenancy. Net soybean and corn revenues received by owner-operators and cash tenants, $p_1 - c$ and $p_2 - hc$, are identical to corresponding net crop revenues generated for the whole farm. Therefore, the farm plan that owner-operators and cash tenants select always maximizes total net farm revenues. Net revenues received by share tenants, $0.6p_1 - c$ and $0.6p_2 - hc$, are less than the net crop revenues received by the entire farm. Where the share tenant's farm plan differs from the owner-operator and cash tenant's farm plan, as is the case within region ABCD, total net farm revenues generated by the share tenant's plan will necessarily be less than the maximum possible. As an example, consider point F. Here, $p_1 = \$230$, $p_2 = \$310$, $c = \$60$, and $h = 2$. The share tenant's plan—200 acres of soybeans—yields total net farm revenues of \$34,000. A maximum possible net revenue of \$38,000 is generated by the owner-operator and cash tenant's plan—200 acres of corn.

Price Mapping Under Risk and Risk Aversion

The introduction of risk and risk aversion into the linear programming model requires adjustments in the objective function. Gross revenues per acre are uncertain, expected values. The operator's reaction to risk is measured as reductions in expected gross soybean and corn revenues per acre. Assume that these reductions, or risk premiums, equal

constant percentages of expected gross revenues. Thus, the risk premium from 1 acre of soybeans equals the given percentage rate multiplied by the expected gross soybean revenue per acre.

With the risk premiums, the objective function becomes:

$$I = rp_1a_1 - ca_1 - r\theta_1p_1a_1 + rp_2a_2 - hca_1 - r\theta_2p_2a_2 \quad (2)$$

Risk premiums per acre of soybeans and corn are θ_1p_1 and θ_2p_2 , respectively. θ_1 and θ_2 are the risk premium coefficients; they indicate the percentages of expected gross soybean and corn revenues deducted to account for risk.

The assumption that risk premiums per acre remain constant as gross crop revenues per acre and total acreages increase may appear restrictive. Risk premiums per acre may in fact be highly correlated with gross crop revenues. Moreover, if covariance between gross soybean and gross corn revenue is nonzero, risk premiums per acre may vary as output ratios vary.

However, if risk premiums per acre are expressed as functions of gross revenues and as functions of output ratios, the objective function will be nonlinear. This nonlinearity will make the analysis more complicated, but little will be gained conceptually. The linear objective function of equation (2) is sufficiently complex to establish the fundamental point of the analysis—namely, that share tenancy may compare even less favorably with owner operation and cash tenancy under risk and risk aversion than under risk neutrality.

Parametric variation of certainty-equivalent net soybean and corn revenues per acre, $rp_1 - c - r\theta_1p_1$ and $rp_2 - hc - r\theta_2p_2$, defines price maps for risk-averse farm operators. Two sets of these maps are derived. In the first, θ_1 is assumed to be greater than θ_2 , indicating that operators consider soybeans riskier than corn. In the second, θ_2 is greater than θ_1 , indicating that corn is riskier than soybeans.

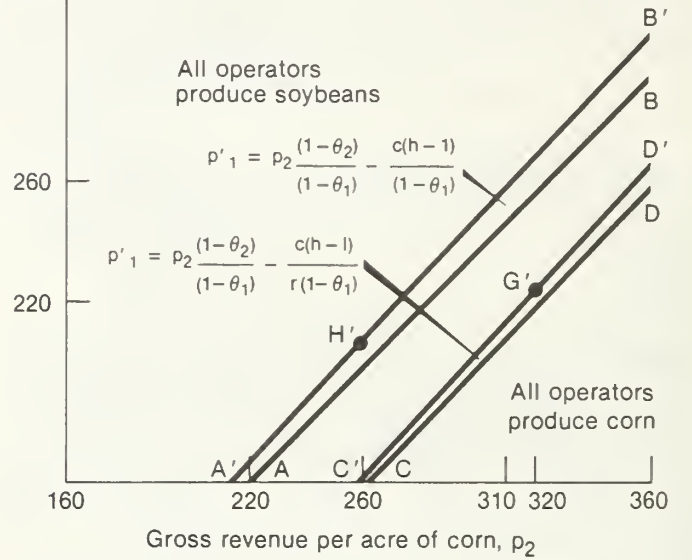
Case 1: $\theta_1 > \theta_2$

Whereas maps AB and CD (fig. 1) equate expected net soybean and corn revenues per acre, price maps for risk-averse farm operators (fig. 2) equate certainty-equivalent net soybean and corn revenues per acre. When $\theta_1 > \theta_2$, equivalency is satisfied on map A'B' for owner-operators and cash tenants and on map C'D' for share tenants. For purposes of comparison, maps AB and CD are also shown in figure 2. To the right of A'B', risk-averse owners and cash tenants receive higher certainty-equivalent net revenues from corn than from soybeans; therefore, they produce only corn. To the left of A'B', they receive more from soybeans than from corn and produce only soybeans. Similarly, to the right (left) of C'D', share tenants receive higher certainty-equivalent revenues from corn (soybeans) than from soybeans (corn); therefore, they produce only corn (soybeans) at all points.

Figure 2

Price Map for Risk-Averse Farmers When Soybeans are Considered Riskier than Corn

Gross revenue per acre of soybeans, p_1



Equality of certainty-equivalent net revenues per acre, $rp_1 - c - r\theta_1p_1 = rp_2 - hc - r\theta_2p_2$, implies that the value of p_1 on A'B' and C'D', hereinafter p'_1 , must be:

$$p'_1 = p_2(1 - \theta_2)/(1 - \theta_1) - c(h - 1)/r(1 - \theta_1) \quad (3)$$

This compares with a value of:

$$p_1 = p_2 - c(h - 1)/r \quad (4)$$

on maps AB and CD. Note that when $\theta_1 > \theta_2$, the slope of A'B' and C'D' ($(1 - \theta_2)/(1 - \theta_1)$) is slightly greater than the slope of AB and CD, which equals 1. However, the vertical intercept of A'B' and C'D' ($-c(h - 1)/r(1 - \theta_1)$) is smaller than that of AB and CD ($-c(h - 1)/r$). It follows that p'_1 in equation (3) is greater than, equal to, or less than p_1 in equation (4) and that A'B' and C'D' therefore lie above, intersect with, or fall below AB and CD, depending on whether:

$$p_2(1 - \theta_2)/(1 - \theta_1) - p_2 \begin{matrix} \geq \\ \leq \end{matrix} c(h - 1)/r(1 - \theta_1) - c(h - 1)/r \quad (5)$$

or, to simplify, on whether:

$$p_2 \begin{matrix} \geq \\ \leq \end{matrix} (\theta_1/\theta_1 - \theta_2)(c(h - 1)/r) \quad (6)$$

The more θ_1 exceeds θ_2 and the smaller h is, the greater is the value of p_2 relative to $(\theta_1/\theta_1 - \theta_2)(c(h-1)/r)$ and the greater are the values of p'_1 on $A'B'$ and $C'D'$ relative to p_1 on AB and CD , respectively.

Suppose the terms in equation (6) are limited to a range of values within which $p_2 > (\theta_1/\theta_1 - \theta_2)(c(h-1)/r)$ and, thus, $p'_1 > p_1$. For example, assume that input costs per acre, c and hc , approximate the recent FED estimates: \$60 for soybeans and \$120 for corn. Also assume that θ_1 exceeds θ_2 significantly. Reasonable values might be: $\theta_1 = 0.04$ per \$1 of expected gross soybean revenue, and $\theta_2 = 0.01$ per \$1 of expected gross corn revenue. It follows from equation (6) that $p'_1 > p_1$ for all $p_2 > (4/3)(60/4)$. Under owner operation and cash tenancy, p'_1 on $A'B'$ will exceed p_1 on AB , and thus $A'B'$ will lie above AB (fig. 2) for all cases where $p_2 > \$80$. Under share tenancy, with $r = 0.6$, p'_1 on $C'D'$ will exceed p_1 on CD , and thus $C'D'$ will lie above CD , for all cases where $p_2 > \$133.33$. Furthermore, given that during the 1974-80 period, the minimum reported gross corn revenue in Illinois was about \$220 per acre, $p'_1 > p_1$ is the only relevant case. Therefore, figure 2 shows only those portions of maps $A'B'$ and $C'D'$ that lie above AB and CD .

For an intuitive explanation of why $A'B'$ and $C'D'$ lie above AB and CD , consider the equality of expected net crop revenues that occurs on AB and CD . Recall that this equality implies that $p_1 = p_2 - c(h-1)/r$. The risk premium per acre of soybeans on AB and CD is $\theta_1 p_1 = \theta_1(p_2 - c(h-1)/r)$. Suppose $c = \$60$, $hc = \$120$, $\theta_1 = 0.04$, and $\theta_2 = 0.01$. Then $\theta_1 p_1$ is greater than $\theta_2 p_2$ (the risk premium per acre of corn) at all points on AB where $p_2 > \$80$ and at all points on CD where $p_2 > \$133.33$. At all such points on AB and CD , certainty-equivalent net revenues are lower for soybeans than for corn. For certainty-equivalent net crop revenues to be equal, as is the case on maps $A'B'$ and $C'D'$, expected net soybean revenue per acre must exceed expected net corn revenue per acre to offset the higher risk premium per acre of soybeans. Thus, for each combination of expected gross soybean and expected gross corn revenues per acre defined by a point on AB or CD , the point on $A'B'$ or $C'D'$ that defines the same gross corn revenue per acre must have a higher gross soybean revenue per acre.

We can also show that p'_1 on $A'B'$ exceeds p_1 on AB by more than the amount that p'_1 on $C'D'$ exceeds p_1 on CD . That is, for any p_2 , the shift from AB to $A'B'$ exceeds the shift from CD to $C'D'$. If p_1 is expressed as in equation (4), then:

$$p'_1 - p_1 = p_2(\theta_1 - \theta_2)/(1 - \theta_1) - \theta_1 c(h-1)/r(1 - \theta_1) \quad (7)$$

which is greater under owner operation and cash tenancy, with $r = 1$, than under share tenancy, with $0 < r < 1$.

Suppose gross corn revenue is \$320 per acre, $\theta_1 = 0.04$, $\theta_2 = 0.01$, $r = 0.6$, and input costs per acre are again \$60 for

soybeans and \$120 for corn. Gross soybean revenue then increases from $p_1 = \$260$ per acre on map AB to $p'_1 = \$267.50$ per acre on map $A'B'$, but only from \$220 per acre on map CD to \$225.83 per acre on map $C'D'$. The risk-induced shift from AB to $A'B'$ is $(\$267.50 - \$260) - (\$225.83 - \$220) = \$1.67$ greater than the shift from CD to $C'D'$.

Thus, when p_2 , c , and hc have reasonable values and when θ_1 exceeds θ_2 by enough to ensure that equal expected net corn and soybean revenues per acre imply higher risk premiums per acre for soybeans than for corn, risk and risk aversion will affect farm planning under owner operation and cash tenancy more than under share tenancy. Share tenancy's standing relative to owner operation and cash tenancy should also improve. Recall that, under risk neutrality, the crop plan chosen by owners and cash tenants is always optimal in the sense that it maximizes total expected net crop revenues for the entire farm. At all points to the left of AB , this optimal plan is 200 acres of soybeans; at all points to the right, it is 200 acres of corn.

To the left of $A'B'$ and to the right of $C'D'$ in figure 2, all risk-averse farmers do in fact choose optimal farm plans. Within region $A'B'AB$, however, only risk-averse share tenants choose the optimal plan—200 acres of soybeans; within $ABC'D'$, only risk-averse owners and cash tenants choose the optimal plan—200 acres of corn.

Within $A'B'AB$, the risk-averse share tenant's plan is more profitable to the farm than the risk-averse owner and cash tenant's plan.⁴ This situation contrasts with the case of risk neutrality (fig. 1). Recall that the risk-neutral share tenant's plan can be equally profitable at best, but can never be more profitable than the owner and cash tenant's plan. Moreover, recall that within region $ABCD$, the risk-neutral share tenant's plan is less profitable to the farm than the risk-neutral owner and cash tenant's plan. Thus, $ABCD$ is analogous to $ABC'D'$, which encompasses an area where the risk-averse share tenant's plan is less profitable to the farm than the risk-averse owner and cash tenant's plan. Observe that $ABC'D'$ is smaller than $ABCD$. That is, the number of soybean-corn revenue combinations that result in a suboptimal share tenant's plan is smaller under risk

⁴ The objective function for risk-averse share tenants who receive proportionate cost share payments from landlords is $0.6(p_1 a_1 - c a_1 - \theta_1 p_1 a_1 + p_2 a_2 - h c a_2 - \theta_2 p_2 a_2)$. Maximization of this function is equivalent to maximization of the risk-averse owner-operator and cash tenant's objective function at all crop revenue combinations. Within $ABA'B'$, the crop plan chosen by risk-averse share tenants who receive proportionate cost share payments as well as by risk-averse owners and cash tenants is less profitable to the farm than the plan chosen by risk-averse share tenants who receive no-cost share payments. This result contradicts the claim of many economists (discussed in footnote 2) that proportionate cost sharing should always increase allocative efficiency under share tenancy.

and risk aversion than under risk neutrality. The probability that the revenue combination in any given year will result in a suboptimal share tenant's plan is also smaller under risk and risk aversion than under risk neutrality.

Therefore, when risk and risk aversion are incorporated into the programming model under the assumption that soybeans are riskier than corn ($\theta_1 > \theta_2$), share tenancy should compare more favorably with owner operation and with cash tenancy than it does in the absence of risk and risk aversion. The probability that the share tenant's plan will be less profitable to the farm than the owner-operator and cash tenant's plan should decline. The probability that the share tenant's plan will be more profitable should increase from zero to some positive value.

Case 2: $\theta_2 > \theta_1$

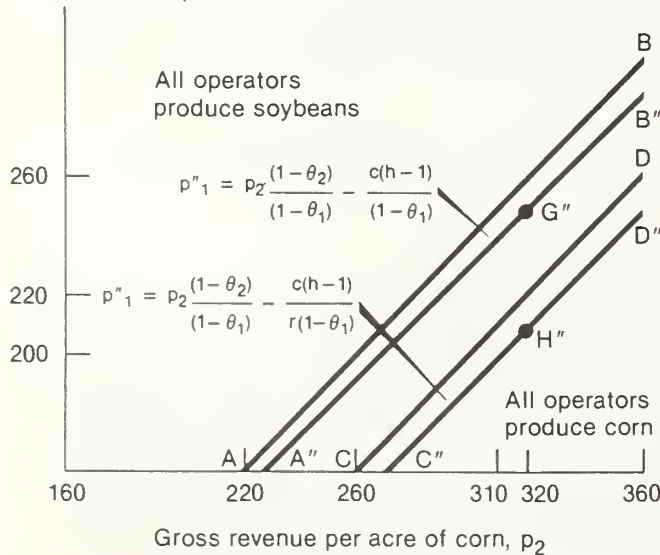
The results are much different if operators consider corn riskier than soybeans. For $\theta_2 > \theta_1$, parametric variation of certainty-equivalent net soybean and corn revenues per acre, $rp_1 - c - r\theta_1 p_1$ and $rp_2 - hc - r\theta_2 p_2$, defines price maps $A''B''$ and $C''D''$ (fig. 3). On these maps, as on maps $A'B'$ and $C'D'$, certainty-equivalent net crop revenues are equal; therefore, the value of p_1 on $A''B''$ and $C''D''$, hereinafter p_1'' , is:

$$p_1'' = p_2(1 - \theta_2)/(1 - \theta_1) - c(h - 1)/r(1 - \theta_1) \quad (8)$$

Figure 3

Price Map for Risk-Averse Farmers When Corn is Considered Riskier than Soybeans

Gross revenue per acre of soybeans, p_1



Note that with $\theta_2 > \theta_1$, the slope of $A''B''$ and $C''D''$, $(1 - \theta_2)/(1 - \theta_1)$, is slightly less than 1, the slope of AB and CD. Furthermore, the vertical intercept of $A''B''$ and $C''D''$ ($-c(h - 1)/r(1 - \theta_1)$) is less than the intercept of AB and CD ($-c(h - 1)/r$). Thus, p_1'' will fall below p_1 , and $A''B''$ and $C''D''$ will fall below AB and CD, for all p_2 . Note also that the more θ_2 exceeds θ_1 and the larger h is, the more p_1 exceeds p_1'' and the greater the gap is between the price maps, for any p_2 . Furthermore, p_1 on CD always exceeds p_1'' on $C''D''$ by more than p_1 on AB exceeds p_1'' on $A''B''$. That is, the shift from CD to $C''D''$ everywhere exceeds the shift from AB to $A''B''$. If p_1 is expressed as in equation (4), then:

$$p_1 - p_1'' = p_2(\theta_2 - \theta_1)/(1 - \theta_1) + \theta_1 c(h - 1)/r(1 - \theta_1) \quad (9)$$

which is higher under share tenancy than under owner operation and cash tenancy.⁵

For example, suppose gross corn revenue again equals \$320 per acre, while $\theta_1 = 0.01$, $\theta_2 = 0.04$, and input costs per acre again equal \$60 for soybeans and \$120 for corn. Gross soybean revenue must then decline from \$220 on map CD to \$209.29 on map $C''D''$, but only from \$260 on map AB to \$249.70 on map $A''B''$. The risk-induced shift from CD to $C''D''$ is \$0.41 greater than the shift from AB to $A''B''$.

Thus, given any values of p_1 , p_2 , c , h , θ_1 , and θ_2 , if $\theta_2 > \theta_1$, risk and risk aversion will always affect farm planning more under share tenancy than under owner operation and cash tenancy. Moreover, the impact of risk and risk aversion on the comparative efficiency analysis, if $\theta_2 > \theta_1$, is quite different than the impact if $\theta_1 > \theta_2$. When $\theta_2 > \theta_1$, we compare $A''B''C''D''$ and ABCD (fig. 3). $A''B''C''D''$ defines combinations of expected gross soybean and corn revenues per acre where risk-averse owner-operators and cash tenants (but not risk-averse share tenants) choose the optimal, profit-maximizing plan—200 acres of corn. Region ABCD defines gross revenue combinations where risk-neutral owners and cash tenants (but not risk-neutral share tenants) choose this same optimal plan. Because the shift from CD to $C''D''$ exceeds the shift from AB to $A''B''$, the number of soybean-corn revenue combinations falling within $A''B''C''D''$ exceeds the number falling within ABCD. That is, the share tenant's farm plan is less profitable to the farm than the owner and cash tenant's plan at more revenue combinations under risk and risk aversion than under risk neutrality. Everything else being equal, the probability that the revenue combination that occurs in any given year will fall within $A''B''C''D''$ will exceed the probability that it will fall within

⁵ Note that for any $\theta_2 > \theta_1$, the excess of $\theta_2 p_2$ over $\theta_1 p_1''$ will decline, and the gap between $A''B''$ and AB and between $C''D''$ and CD will narrow as p_1' and p_2 decline.

ABCD. Thus, the probability that the share tenant will actually choose a suboptimal plan will also be higher under risk and risk aversion than under risk neutrality.⁶

Comparison of Crop Risks

This analysis has shown that the impact of risk and risk aversion on the comparative efficiency analysis presented by the programming model depends on the values of the risk premium coefficients assigned to the alternative crops. This result raises the question of which of the two sets of values assumed in this article better represents the attitudes of farm operators toward the riskiness of soybeans and corn.

One answer is suggested by estimates of the variability of gross soybean and corn revenue that can be derived from recent average annual yield and price data provided by USDA's Crop Reporting Board (37, 38, 39, 40, 41, 42, 43, 44). Let us take Illinois as an example. A sample of seven gross corn and soybean revenues per acre of Illinois cropland from the 1974-80 period indicates recent variations. The sample standard deviation was approximately \$39 per acre for corn and \$33 per acre for soybeans. Corn, therefore, exhibited greater variability than soybeans. Thus, a typical Illinois farm operator would likely consider corn as the riskier crop. The assumption that $\theta_2 > \theta_1$ appears to be more appropriate than the assumption that $\theta_1 > \theta_2$.

If so, the conclusion for case 2 applies. Share tenancy will likely compare less favorably with owner operation and cash tenancy under risk and risk aversion than under risk neutrality.

Conclusion

Previous analyses of the comparative efficiency of different farm tenancies under risk and risk aversion have relied exclusively on single-input, single-output models. When risk and risk aversion are introduced into these models, the efficiency of share tenancy relative to owner operation and cash tenancy always increases.

When the comparative efficiency analysis is extended to multiproduct programming models, however, the results can be quite different. Under risk neutrality, share tenancy compares unfavorably with owner operation and cash tenancy. Within a certain range of crop revenue combinations, share tenants

choose a crop plan which is less profitable to the farm as a whole than the plan owner-operators and cash tenants choose.

To determine the impact of risk and risk aversion on this comparative efficiency analysis, we examined the implications of two different assumptions concerning reactions of farm operators to income variability. Case 1 assumed that farm operators assign a higher risk premium to \$1 of gross soybean revenue than they assign to \$1 of gross corn revenue, even though the variable production cost per acre of corn is higher than the cost of soybeans. The number of soybean-corn revenue combinations that call for a suboptimal share tenant's plan is then lower under risk and risk aversion than under risk neutrality. Some revenue combinations even result in a risk-averse share tenant's plan that is more profitable than the risk-averse owner-operator and cash tenant's plan. The probability that the share tenant's plan will be less profitable than the owner and cash tenant's plan is unequivocally lower under risk and risk aversion than under risk neutrality.

In case 2, risk and risk aversion were introduced under the assumption that the risk premium per \$1 of gross corn revenue is higher than the risk premium per \$1 of gross soybean revenue. The number of soybean-corn revenue combinations that result in a suboptimal share tenant's plan is now greater under risk and risk aversion than under risk neutrality. Therefore, the probability that the share tenant's plan in any year will be less profitable than the owner and cash tenant's plan is also greater under risk and risk aversion than under risk neutrality.

An empirical review of crop and yield data for Illinois suggests that the second of the two assumptions—that is, that the risk premium per \$1 of corn is greater than the risk premium per \$1 of soybeans—more accurately represents Illinois agriculture. As the analysis in case 2 appears unfavorable to share tenancy, a contrast between the single-input, single-output model and the multiproduct programming model becomes obvious. Although the presence of risk and risk aversion always improves share tenancy's standing relative to owner operation and cash tenancy under the first model, it probably weakened share tenancy's relative standing under the second model, at least for Illinois agriculture during the late seventies.

Multiproduct models may generate different results for other regions and other time periods. Where the analysis in case 1 applies, share tenancy will compare more favorably with owner operation and cash tenancy under risk and risk aversion than under risk neutrality. When risk and risk aversion are relevant, the comparative efficiency of alternative farm tenancies can no longer be determined by theoretical considerations alone. Comparative efficiency can be determined only after one has assigned observed or reasonable values to the model parameters.

⁶ It is conceivable, if improbable, that despite the greater number of revenue combinations within A'B'C'D' than within ABCD, the probability that the combination in any given year will fall within ABCD still exceeds the probability that it will fall within A'B'C'D'. The probability that the share tenant will choose an inefficient plan will then be lower under risk and risk aversion than under risk neutrality. The introduction of risk and risk aversion into the model will offset, rather than augment, the distortive effect of product-sharing on the relative efficiency of share tenancy.

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Prefiltering and Causality Tests

By Mike Belongia and David A. Dickey*

Abstract

If data series are not filtered properly prior to the construction of a test of causality, the resulting test statistics are invalid. This article describes a general approach to data filtering based on the estimation of autoregressive-moving-average models and on specific tests for the identification of white noise processes. For selected examples, traditional approaches to filtering do not perform as well as the general method proposed.

Keywords

Causality tests, data filtering, econometric exogeneity

Statistical techniques developed for testing whether the behavior of one variable causes a subsequent change in the value of another variable have been widely used by economists in recent years. The tests are often used to determine whether one variable can be treated as exogenous with respect to another. The tests are also used as a form of pretest estimation in determining whether a regressor makes a significant contribution to the explanation of the variation in a dependent variable.

Although causality tests have become more widely used in recent years, many practitioners disagree over the validity of the results of these techniques.¹ This article reviews some empirical issues relevant to the use of causality tests and the necessary documentation of the procedures employed. Because the type of test employed in much applied research depends crucially on proper filtering of the data prior to testing, we will focus on problems that are likely to occur if economic data are transformed by the filter suggested by Sims, but the results are not checked against some white noise test.

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¹ The problem of detecting causal relations can be especially acute in bivariate models if the variables are likely to share a common relationship with a third variable; this problem is discussed at length in (8), (9), and several other papers. (Note: Italicized numbers in parentheses refer to items in the References at the end of this article). Schwert (11) has argued that Box-Jenkins models may not be causal-preserving so that the use of autoregressive-moving-average (ARMA) model residuals in causality tests makes those tests subject to a potential errors-in-variables problem. That is, "if the original variables are measured with random errors, causality tests based on the estimated innovations series could fail to detect relationships that would be detected using the untransformed data." A recent review of alternative testing procedures and their limitations is summarized in (1).

We first outline the intuition supporting causality tests and discuss criticisms raised by Feige and Pearce regarding the testing procedures used in much of the applied literature (4). We then illustrate the potential problems associated with applying the Sims filter to economic data with examples using economic time series frequently employed in studies of inflation and price change. We discuss an alternative approach to data filtering and diagnostic checks for white noise that uses an autoregressive-moving-average (ARMA) transformation. Finally, we make suggestions regarding the future use of causality tests.

Causality Tests: An Overview

Granger (6) defines econometric causality as follows: " Y_t is causing X_t if we are better able to predict X_t using all available information than if information apart from Y_t has been used." Or more simply, Y_t causes X_t in the econometric sense if and only if one can predict X_t better by using past values of Y_t and X_t rather than by basing the prediction on the past history of X_t alone. Sources other than the original papers by Granger and Sims provide details on the mechanics of applying such a test (see (2)). The general idea is to estimate regressions with and without the additional information contained in Y . From these regressions one can construct a joint F-test on the significance of the coefficients associated with future values of Y . A significant F value would then suggest the presence of a causal relationship between X and Y . To check for feedback—or whether causation also runs from Y to X —one can repeat the test by regressing Y on past and future values of X . If neither joint F-test on the coefficients associated with future values of X or Y is significant, one can conclude that the two variables are unrelated.

The importance of filtering the data prior to estimating the regression models and calculating the joint F-statistic is related to the spurious regression problem described by

Granger and Newbold. Data filtering is intended to remove the serial correlation present in most time series. If filtered properly, each resulting time series used in a test of causality will be a white noise process so that any significant relationships represented by the joint F-test will be based on actual, systematic relationships between the two series instead of on a spurious relationship caused by the common serial correlation.

Adequate data filtering is essential to the validity of causality tests because the failure to remove serial correlation from the data will bias the estimates of coefficient variances. Because most time series share a similar pattern of serial correlation, the variances are likely to be biased downward. This downward bias will result in artificially large t statistics associated with individual coefficients and in a correspondingly large joint F-statistic for any causality test. Failure to filter the data adequately may leave some common serial correlation in the data that will bias the test statistics upward and possibly suggest a significant causal relationship where, in fact, none exists.

Data Filtering

The problem of serial correlation might be turned into a benefit if a common pattern among time series makes it possible for one filter to transform successfully most economic series to the white noise processes required by the testing procedures. Such a filter would standardize the mechanics of testing procedures and let researchers conduct their tests of interest after a routine data transformation. Although no universal filter has been found, some applied researchers apparently believe that such a filter exists. This belief creates a problem for those who apply and interpret the results of causality tests.

Belief in a universal data filter is probably the result of a statement that Sims made in his original article.² In a study which investigated the causal relationships between gross national product (GNP) and different measures of the money supply, Sims advocated filtering the data by the following transformation:

$$Z = \ln X_t - 1.5 \ln X_{t-1} + 0.5625 \ln X_{t-2}$$

which is an expansion of $(1 - KL)^2$ where L is the lag operator and $K = 0.75$. He said this filter "approximately flattens the spectral density function of most economic time series, and the hope was that regression residuals would be very nearly white noise with this prefiltering" (12). As we shall see later, the spectrum of a white noise process will be flat

and its plot can be used as one diagnostic check for the adequate filtering of data.

The (unintended) result of Sims' statement has been for researchers to apply his filter to a wide variety of economic time series *without* subsequent checking for whether this transformation actually has created a new white noise process. As the following examples will illustrate, many common economic time series are not transformed to white noise by Sims' filter. This result is not wholly unexpected in view of the numerous and volatile shocks represented in the economic data since 1972 when his article was published. But, unless the data are transformed to white noise processes, the results of a causality test are invalid.

Sims' Filter Applied to Some Common Economic Series

Prior to the description of an alternative approach to data filtering, it may be helpful to illustrate the potential problems associated with not testing filtered data to determine whether the transformation has reduced a series to a white noise process. For this purpose, the following variables have been chosen: the narrowly defined money stock (M1), average wage rates for the manufacturing sector (W), and the GNP deflator (DEF).³ Each series is monthly from January 1961 through December 1977. Table 1 presents descriptive statistics for each series—prior to transformation and after transformation by the Sims filter.

The spectra for these series are plotted in figures 1-3. The area under a spectral plot is the variance of the data series. Because we have moved from the time domain to the frequency domain, the frequencies represented on the horizontal axis are measured from $-\Pi$ to Π radians; however, because the plot is symmetric about zero, only the area from 0 to Π is shown.⁴

Visually, the spectrum identifies any spikes in the plot associated with particular frequencies between 0 and Π . The presence of a spike in the spectral density at a particular frequency suggests that a relatively larger share of the series' variance is explained by that frequency. If the series is a white noise process, the spectrum should not contain any spikes because no one frequency would contribute more than any other to the explanation of variance. A visual check of the spectrum can be supported by the Kolmogorov-Smirnov and Fisher tests for white noise; these tests are described in several texts (see (5)).

³ The monthly GNP deflator series was provided by Data Resources, Inc.

⁴ The spectra for these series were estimated with triangular weights of the form 1-2-3-2-1 and 1-2-3-4-3-2-1; the results did not vary with the choice of weighting scheme. Issues associated with the choice of lag window and the problem of leakage are discussed in most time-series texts.

² A similar problem results if the data are expressed as first differences, a data transformation suggested by Box and Jenkins as a means of detrending a time series (3).

Table 1—Descriptive statistics for variables¹

Variable	Mean	Standard deviation	Minimum	Maximum
<i>In levels</i>				
DEF	0.938	0.230	0.688	1.453
M1	216.918	57.168	144.000	348.200
W	3.472	1.022	2.280	5.880
e_t	.000	.001	-.004	.003
<i>Transformed by Sims' filter</i>				
DEF	-.005	.015	-.025	.026
M1	.335	.024	.284	.390
W	.364	.019	.333	.404

¹ Number of observations = 204.

The Sims filter did not transform any of the three series to white noise. The spectral plots display a result common to most economic series—that is, a large spike in the spectrum at low frequencies near the origin. A spectrum of this shape suggests that much of the variance of the series can be explained by a strong trend in the data. However, the intention of the filtering was to remove elements of trend so that the transformed series would be a white noise process. The plots in figures 1-3, which result from filtering common aggregate series by Sims' method, show clearly that this method is not appropriate for these data. The conclusions suggested by the spectral plots are supported further by the white noise tests

reported in table 2. For each series, the test statistic rejects the null hypothesis that the chosen filtering technique produced a white noise process. Causality tests based on these data could yield invalid test statistics.

Figure 2

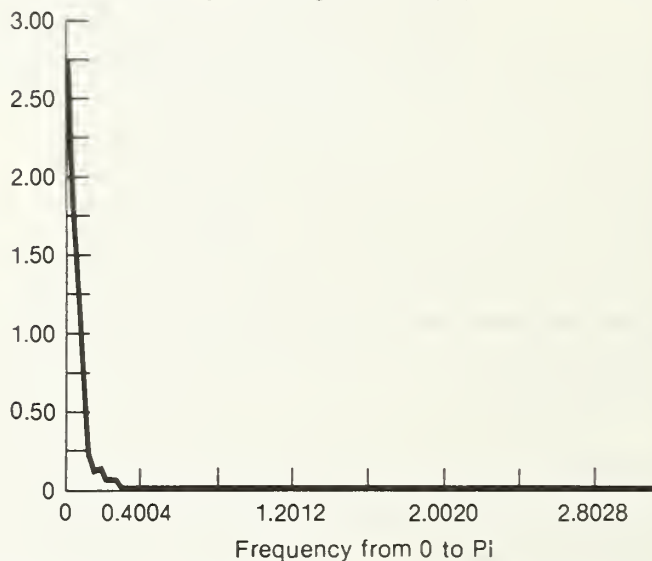
Spectral Density for Wage Rates (W)Vertical scale entries are multiplied by 10^3 .

Figure 1

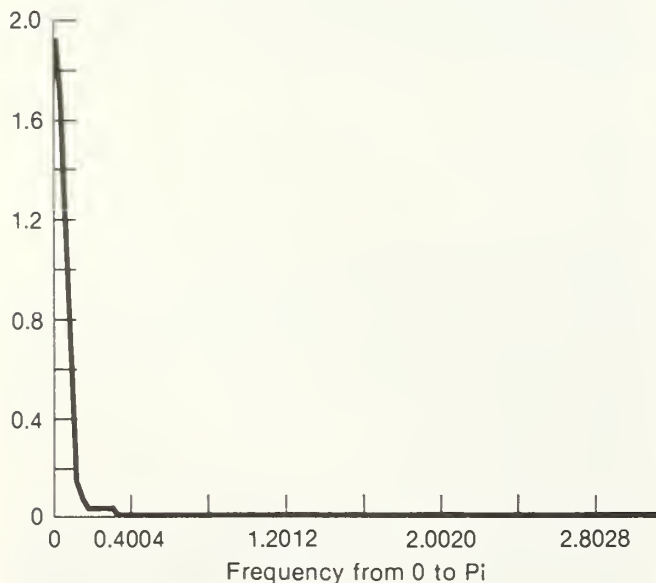
Spectral Density for GNP Deflator (DEF)Vertical scale entries are multiplied by 10^3 .

Figure 3

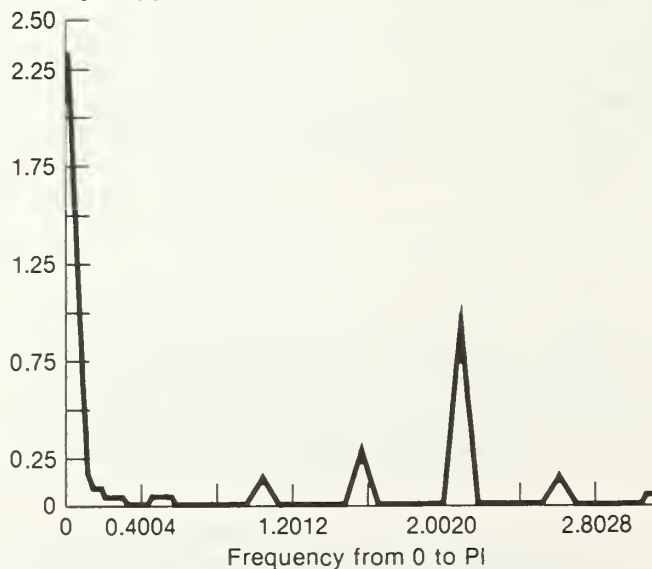
Spectral Density for Narrowly-Defined Money Supply (M1)Vertical scale entries are multiplied by 10^3 .

Table 2—Estimated values for the Fisher Kappa statistic

Variable	Estimated Fisher Kappa ¹
DEF	64.31
W	59.89
MI	30.76
e_t	5.89

¹ All estimated values are to be compared with a critical value of 7.38. Critical values are provided by (5, p. 284).

An Alternative Approach to Filtering

One may search for an appropriate filter by changing the value of K in Sims' transformation or by attempting similar *ad hoc* data manipulations. We will now present an alternative approach to filter selection and use one of the previous series to illustrate its practical application. Although this technique involves some preliminary data analysis, it can provide a reasonable guide to the identification of an appropriate data transformation.

The suggested alternative to filtering uses the estimated coefficients of an ARMA model to transform each data series. This approach requires identifying each series based on an analysis of its autocorrelation and partial autocorrelation functions, estimating the identified model, and finally, getting the model's estimated residuals. If the fitted model is the appropriate time series representation of the data, its residuals will be white noise. However, identification of the most appropriate time series representation of a variable is not always a simple task: "It is important to realize the specification of an ARMA model is an art, rather than a science" (10).

The residuals from each model can be tested against a null hypothesis of white noise in the same manner already described. Once a vector of white noise residuals is created for each series, the actual causality tests can be run by use of these residual vectors. The causality test will regress one vector of residuals on past and future values of the other residual vector. A joint F-test on the significance of the coefficients associated with the future values will then indicate if, after proper filtering, one series still contributes to the explanation of the variation in the other.

To illustrate, the GNP deflator (DEF) will be identified and then estimated by an ARMA model with the residuals from the estimated model subsequently tested for white noise. If we were to test for a causal relationship between DEF and another variable, the same steps would be followed, thereby creating a residual vector for that series. We could then perform the actual causality tests by regressing the residual vectors on each other as described earlier. However, our purpose here is only to illustrate an alternative approach to data filtering.

The first step involves the identification of an ARMA model which will represent the process that generates values for the series. Procedures described by Box and Jenkins and found in a variety of time series texts suggest that one can identify a model by analyzing the autocorrelation and partial correlation functions of the series. The plots of these functions indicate how many autoregressive and moving average terms to include in the model. For DEF, these plots suggest that a third-order autoregressive model will adequately represent the process which generates values for DEF.⁵ That is, we have identified a model of the form:

$$DEF_t = \alpha + \beta_1 DEF_{t-1} + \beta_2 DEF_{t-2} + \beta_3 DEF_{t-3} + e_t$$

to represent the DEF series.

With estimated values for the α and β_i terms, a residual vector can be created by simple manipulation:

$$e_t = DEF_t - \alpha - \beta_1 DEF_{t-1} - \beta_2 DEF_{t-2} - \beta_3 DEF_{t-3}$$

This estimated residual vector is the filtered series which would be used in a causality test. If the estimated ARMA model is the correct model for the DEF series, the e_t vector should be white noise.

To test the residuals as a white noise process, one can employ plots of the series' spectral density and white noise tests. A flat spectral density indicates a white noise series because no particular frequency between 0 and Π radians makes a relatively larger contribution to the series' variance than any other. This visual check can be supported by a white noise test provided by Fisher (see (5)).

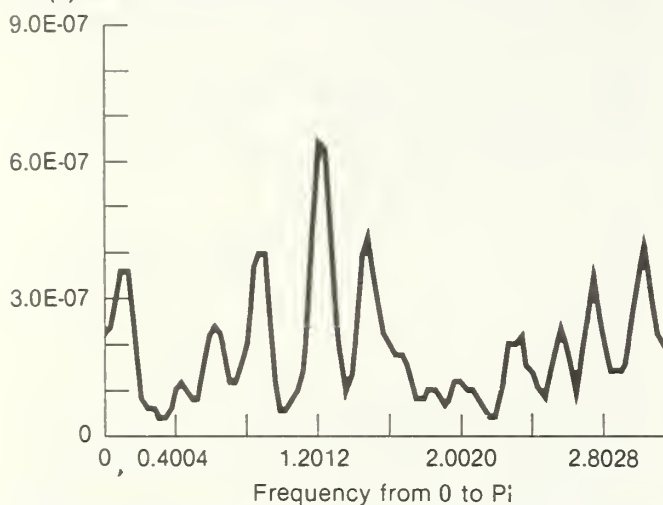
Figure 4 shows the spectral density for e_t . Table 2 shows the Fisher Kappa statistics for all variables. The plot reveals no particular pattern in the relationship between frequency and the height of the spectral density. This flat spectrum would suggest that e_t is a white noise process. This result is supported by a test statistic of 5.89, which is less than the critical value of 7.38 for 100 degrees of freedom.⁶ From these test results, one can conclude that the residual vector from the estimated ARMA model is a white noise series. Thus, the e_t vector is an appropriately filtered representation of DEF that could be used in a causality test.

⁵ One should note that, although a third-order autoregressive model was found to be appropriate for the DEF series, this same model may not appropriately represent other series. In fact, the DEF series could most likely be represented and transformed to white noise by something other than an AR(3) model, a possibility discussed in most time series texts. However, the point remains that filtering data by using the coefficients of an ARMA model requires that each time series be analyzed individually rather than be transformed by some common manipulation.

⁶ The Fisher Kappa does not follow a standard F distribution. Fuller provides the correct critical values against which the estimated test statistic should be compared (5, p. 284).

Figure 4

Spectral Density for Residuals From the AR(3) Model Estimated for the GNP Deflator



Conclusions

One of the requirements of causality tests is that the data series be filtered to create white noise vectors from the original series values. Sims suggested a filter which he apparently believed would adequately transform most economic series—at least prior to 1972. We have demonstrated, however, that this filter does not properly filter a variety of common economic series, thus invalidating their use in causality tests.

A suggested alternative to filtering involves estimating an ARMA model. The residuals from the ARMA model are the filtered version of the data series. The spectral plot for the residual vector and tests for white noise can be used to indicate whether the results are, in fact, a properly filtered white noise process. Results based on the GNP deflator (DEF) indicate that Sims' filter does not produce a white noise series but that the residuals from an ARMA model are white noise and can, therefore, be used correctly as one variable in a causality test.

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Input Demand Formulations and Duality Theory

By Thomas J. Lutton*

Abstract

It is difficult to obtain closed-form solutions for input demand equations that are conducive to econometric estimation and policy analysis when the production function is unknown. Developments in the theory of duality facilitate obtaining such equations. This article provides three examples of dual approaches to neoclassical constrained optimization. One can obtain input demand equations and all comparative statics effects from the first and second derivatives of these dual functions.

Keywords

Demand, duality, cost function, indirect production, function, profit function

Policy analysis of input demand frequently requires the systematic assessment and projection of the effects of an input price change on input demand, costs of production, and output. When comparative, static, competitive equilibrium assumptions are appropriate and when the typical producer's production possibility set is known, neoclassical production theory provides the analyst with an analytical framework for evaluating such effects.

Unfortunately, the policy analyst, unlike the producer, seldom has a quantitative *a priori* characterization of the production possibility set or of the production function that determines input demand. Assuming that the analyst chooses a neoclassical parametric approach to input demand formulation, he or she typically assumes some approximate functional form for technology, uses this form in a constrained optimization, and derives the input demand equations from the first- and second-order conditions. The analyst attempts to use observations consisting of input and output quantity and price data directly or indirectly to estimate an input demand function.

Advances in duality theory allow us to formulate a system of input demand functions which can be used to examine such issues as input price sensitivity, factor substitution, and output effects. Cost, indirect production, and profit functions are particularly useful for specifying and estimating input demand functions and for analyzing comparative statics. When the production function is unknown, these input demand functions are more easily derived with the dual approach

than with the traditional, constrained optimization approaches. When the production function is unknown, the dual approach circumvents both the statistical problems associated with direct estimation of production functions and the computational difficulties in obtaining input demand functions for estimation.

The analyst pursuing a neoclassical parametric approach has three choices: to econometrically estimate the assumed production function and use this functional form in a constrained optimization to derive input demand schedules, to estimate the first-order conditions assuming some functional form for the marginal product of each input, or to estimate the reduced-form input demand system which explicitly includes both first- and second-order conditions.¹

In the first approach, one can estimate the parameters of the production function with single-equation estimators, ignoring factor prices. There is, however, a statistical problem with this approach. Such estimates of the parameters of the production function will be biased if inputs and outputs are jointly determined. Varian provides an example of bias in the direct estimation of a production function for corn (34). If the regressors (input quantities) in a production function are correlated with the error term, ordinary-least-squares estimates will be biased. In addition, if the production function is unknown and is approximated by a truncated expansion such as a Taylor series to avoid predetermining the elasticities of substitution, the number of parameters and arguments (squares, cross products, and so forth) in a single equation for more than three inputs becomes extremely large,

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¹ Nonparametric approaches to input demand are contained in (1, 11, 16). Note: Italicized numbers in parentheses refer to items in the References at the end of this article.

complicating the estimation and possibly reducing the analyst's confidence in the parameter estimates.²

The second option is to estimate the first-order conditions with factor prices as dependent variables and input quantities as exogenous variables. If the producer is assumed to be a price taker, the causality is reversed in the econometric model. Furthermore, if input quantities are jointly determined, multicollinearity problems may be severe.

The third option is the preferred method for analyzing the systems effects of an input price change, data permitting and assumptions warranted. However, a difficult methodological problem arises with this approach. It is generally possible to obtain the reduced-form or closed-form solution only for the simpler functional forms of the production function. This difficulty may prevent formulation of the input demand equations derived from any second-order approximation to production functions which are more complex. Such flexible approximations are increasingly used for econometric estimation and for subsequent, comparative static analysis because they avoid unnecessary restrictions on the elasticities of substitution.

The theory of production duality has its roots in Hotelling (18), Hicks (17), Samuelson (30), Shephard (31), McFadden (23), Uzawa (33), and Diewert (9). The dual approach remained a didactic tool until Nerlove (27) used a cost function to derive and estimate reduced form input demand equations. Diewert (9), Fuss (12), Christensen, Jorgenson, and Lau (5,6), Denny (7), Norman and Russell (28), and Berndt and Kahled (3) contributed flexible functional forms for dual relationships.

The emphasis here is on input demand formulation for empirical analysis. I derive input demands from three alternative objective functions for the competitive producer: output-constrained cost minimization, cost-constrained output maximization, and profit maximization. Each yields a derived demand formulation with different exogenous arguments.

The producer is assumed to be a price taker in both input and output markets. Here I briefly review comparative statics using the neoclassical approach and compare this approach with dual formulations when the production function is unknown.

² Consider a translogarithmic production function where output (Q) is expressed as a function of input quantities $X = (X_1, X_2, \dots, X_n)$:

$$\ln Q = \alpha_0 + \sum_j B_j \ln X_j + \frac{1}{2} \sum_i \sum_j B_{ij} \ln X_i \ln X_j$$

where α_0 , B_i , B_{ij} are constants for $i, j = 1, 2, \dots, n$. If $B_{ij} = B_{ji}$ and $n = 5$, 21 coefficients must be estimated.

Neoclassical Comparative Statics

As Mundlak (26) has demonstrated, it is possible to cast all three problems in a common framework. A competitive firm is assumed to have a twice differentiable production function:

$$Q = F(X) \quad (1)$$

where Q is an output and X a vector of input quantities; $X = (X_1, X_2, \dots, X_n)$.

The first order conditions for each of the three constrained optimization problems can be written as:

$$F_i = \lambda P_i, \quad i = 1, 2, \dots, n \quad (2)$$

where $F_i = \partial F(X) / \partial X_i$, P_i is the price of factor i , and λ is a Lagrangian multiplier. Following Mundlak, the displacement from equilibrium for these three cases can be represented as:

$$\begin{bmatrix} -\hat{\lambda} \\ dX_1 \\ dX_2 \\ \cdot \\ \cdot \\ dX_n \end{bmatrix} = \lambda \begin{bmatrix} 0 & F_1 & F_2 & \dots & F_n \\ F_1 & F_{11} & F_{12} & \dots & F_{1n} \\ F_2 & F_{21} & F_{22} & \dots & F_{2n} \\ \cdot & \cdot & \cdot & \dots & \cdot \\ F_n & F_{n1} & F_{n2} & \dots & F_{nn} \end{bmatrix}^{-1} \begin{bmatrix} dZ \\ dP_1 \\ dP_2 \\ \cdot \\ \cdot \\ dP_n \end{bmatrix}$$

$$= \begin{bmatrix} K_{oo} & K_{oj} \\ \cdot & \cdot \\ K_{io} & K_{ij} \end{bmatrix} \begin{bmatrix} dZ \\ dP_1 \\ dP_2 \\ \cdot \\ \cdot \\ dP_n \end{bmatrix} \quad (3)$$

where $\hat{\lambda} = d \log \lambda$. The three cases differ in the measurement of dZ , the value of the constraint at equilibrium. These values are respectively:

$$(a) \text{ Cost minimization } (dZ = \frac{1}{\lambda} dQ = 0) \quad (4a)$$

$$(b) \text{ Output maximization } (dZ = dC - \sum X_i dP_i = \sum X_i dP_i) \quad (4b)$$

$$(c) \text{ Profit maximization } (\hat{\lambda} = 0, \\ dZ = - (1/K_{oo}) [K_{oj}] [dP]) \quad (4c)$$

Utilizing equations (3) and (4), Mundlak shows that the substitution effect in output-constrained cost minimization is identical to that in cost-constrained output maximization and profit maximization. Similarly, cost-constrained output maximization yields an output effect which is identical to that in a profit-maximizing adjustment. The distinguishing feature among the derived demand formulations is what remains fixed: output (Q), expenditures (C), or output price (P_Q). Using Mundlak's notation:

Substitution Effect (dQ = 0):

$$\left. \frac{\partial X_i}{\partial P_j} \right|_{dQ=0} = K_{ij} \quad (5a)$$

Substitution and Output Effect (dC = 0):

$$\left. \frac{\partial X_i}{\partial P_j} \right|_{dQ=0} = K_{ij} - \left. \frac{\partial C}{\partial P_j} \right|_{dQ=0} \left. \frac{\partial X_i}{\partial C} \right|_{dP=0} \quad (5b)$$

Substitution, Output, and Maximizing Effect (dP_Q = 0):

$$\left. \frac{\partial X_i}{\partial P_j} \right|_{d\hat{\lambda}=0} = K_{ij} - \left. \frac{\partial C}{\partial P_j} \right|_{dQ=0} \left. \frac{\partial X_i}{\partial C} \right|_{dP=0} \\ + \left. \frac{\partial C}{\partial P_j} \right|_{d\hat{\lambda}=0} \left. \frac{\partial X_i}{\partial C} \right|_{dP=0} \quad (5c)$$

The desirability of the closed-form solution is apparent, because Mundlak has demonstrated that the substitution, output, and maximizing effects of an input price change may be obtained with partial derivatives of the closed-form solution, thus making inversion of the bordered Hessian unnecessary. Except for restrictive functional forms, the input demand formulation is difficult when the researcher starts with a neoclassical approach as described by Mundlak.

Cost Function: A Dual Approach

Fortunately, one need not specify the production function to derive the input demands. Consider the cost-minimizing input demand. Using the duality principles of Shephard (31) and Samuelson (30), Diewert (9) demonstrated how to obtain the cost-minimizing input demand by postulating

the existence of a generalized cost function, dual to F(X), and by employing Shephard's lemma, how to derive the optimal input demand, X_i^* .³ The cost function C(Q, P_X), where costs are a function of output and input prices, is defined by:

$$\text{Min } C = C(Q, P) = \text{Min } (\sum P_i X_i \mid F(X) \geq Q \text{ and } X, P \geq 0) \\ = P_i X_i^*(Q, P) \quad (6)$$

where $X_i^*(Q, P)$ is the closed form solution for input i.

Shephard, Samuelson, and McFadden have established the duality between cost and production functions. Given one of these functions, together with the usual convexity, homogeneity, and differentiability assumptions made with respect to it, the other function is uniquely determined (32). For a cost function C(Q, P) to be dual to a production function, it must satisfy the following properties:

- Positive for all Q and P,
- Nondecreasing in Q and P,
- Linearly homogeneous in P, and
- At least quasi-concave in P.

The production function dual to the above cost function will be of the classical type: nonnegative for nonnegative X, nondecreasing in X, and at least quasi-concave in input quantities.

An important advantage of the dual approach to production analysis is that it often affords a simpler analytical procedure than the traditional approach. If a well-behaved C(Q, P) exists for a producer, Shephard's lemma states that the output-constrained input demand for the ith input is equal to the partial derivative of C(Q, P) with respect to P_i:

$$X_i^* = \frac{\partial C(Q, P)}{\partial P_i} \quad (7)$$

Moreover:

$$\left. \frac{\partial X_i}{\partial P_j} \right|_{dQ=0} = \frac{\partial^2 C(Q, P_1, \dots, P_n)}{\partial P_i \partial P_j} = K_{ij}$$

An example of a Cobb-Douglas type cost function with its dual production function is provided. Write the cost function as:

³We use the notation X_i^* , $\bar{X}_i^*$, and $\hat{X}_i^*$ to designate the optimal input quantities associated with the cost-minimizing, output-maximizing, and profit-maximizing objective functions, respectively. We use different notations for each input demand to emphasize that the comparative statics effects of an input price change are generally different for each.

$$C(Q, P) = b_0 P_1^{b_1} P_2^{b_2} Q \quad (8)$$

where b_0 , b_1 , and b_2 are positive constants. $C(Q, P)$ is thus assumed to be increasing in P_1 , P_2 , and Q . If we assume constant returns to scale in production, $b_1 + b_2 = 1$. The optimal demand for X_1 would then be expressed as:

$$X_1^* = b_1 b_0 P_1^{b_1-1} P_2^{b_2} Q \quad (9)$$

A Cobb-Douglas production function dual to the Cobb-Douglas cost function (8) may be written:

$$Q = \frac{b_1^{-b_1} b_2^{-b_2} X_1^{b_1} X_2^{b_2}}{b_0} \quad (10)$$

The constrained Lagrangian cost-minimizing input quantity X_2 , given equation (10), is identical to that in (9):

$$X_1^* = b_1 b_0 P_1^{b_1-1} P_2^{b_2} Q \quad (11)$$

Indirect Production Function: A Dual Approach

Econometric estimates of input demands derived from the cost function have been applied to agriculture by Duncan and Webb (10), Mensah and Miranowski (23), and Reisner and LeBlanc (22). The cost function is appropriate for investigating the factor-substitution effects of an input price change that is conditional upon output level. If, however, the producer is an output maximizer with a budget constraint, the cost function is inappropriate for derived demand formulation. Moreover, estimation of equation (7) will produce biased parameter estimates, as output is likely to be correlated with the error term. The objective of the analyst may be to evaluate the expected effect on output of an increase in an input price as well as the possible decrease in the input per unit of output. The substitution and output effects in equation (5b) may be difficult to obtain in the constrained Lagrangian framework if $F(X)$ is sufficiently flexible and if inversion of the bordered Hessian is required.

However, one can obtain the combined substitution and output effects by using the duality relationships without taking the inverse of a matrix. Hanoch (15) shows that it is generally possible to postulate the existence of an indirect or polar production function dual to a given production function.

Let us define the indirect production function as:

$$Q = Q(C, P) = \text{Max } F(X) \text{ s.t. } \sum_{i=1}^n P_i \bar{X}_i^* = C \quad (12)$$

where $\bar{X}_j^*$ is the optimal input quantity maximizing Q and satisfying the budget constraint. $Q(C, P)$ is referred to as the

indirect production function in which output is a function of input prices and costs of production. In the same fashion that $C(Q, P)$ may be uniquely derived from $F(X)$ in output-constrained cost minimization, $Q(C, P)$ may be uniquely derived from $F(X)$ in cost-constrained output maximization. For this derivation to be possible, the function $Q(C, P)$ must possess the following properties:

- Decreasing in P and increasing in C ,
- Quasi-convex in P , and
- Homogeneous of degree zero in P and C .

It is possible to derive the optimal demand for an input (using Roy's lemma) as:

$$\bar{X}_i^* = \bar{X}_i^*(P, C) = - \frac{\partial Q(C, P)}{\partial P_i} / \frac{\partial Q(C, P)}{\partial C} \quad (13)$$

where $\bar{X}_i^*$ is the optimal input demand, given the cost-constrained, output-maximization objective function. For any production function $F(X)$, Hanoch proves the unique existence of the indirect production function $Q(C, P_X)$. Similarly, given a well-behaved indirect production function, it is possible to show that a unique production function $F(X)$ exists. Thus, it is possible to specify $Q(C, P_X)$ and estimate input demands directly from the indirect production function. Note that $\bar{X}_i^*$ is defined in terms of cost and input prices, not in terms of output quantity and input prices. Evaluating the total substitution and output effect of an input price change yields:

$$\left. \frac{\partial \bar{X}_i^*}{\partial P_j} \right|_{dC=0} = \frac{\partial \bar{X}_i^*(C, P)}{\partial P_j} \quad (14)$$

Thus, it is possible to anticipate the change in the demand for an input i as the derivative of the demand function $\bar{X}_i^*(C, P_X)$. It is preferable to employ $Q(C, P)$ rather than $F(X)$ when one derives equation (13) and observes changes in input demand under cost-constrained output maximization if the form of the production function $F(X)$ is unknown.

To separate the substitution and output effects in equation (14), observe that if the form of $Q(C, P)$ is known, it is possible then to write $C = C(Q, P)$. Using Mundlak's results, one can show that:

$$\left. \frac{\partial \bar{X}_i^*}{\partial P_j} \right|_{dC=0} = \frac{\partial^2 C(P, Q)}{\partial P_j^2} \bigg|_{dQ=0} - \frac{\partial C(P, Q)}{\partial P_i} \bigg|_{dQ=0} \frac{\partial Q(P, C)}{\partial P_j} \bigg|_{dC=0} \quad (15)$$

where the substitution and output effects, respectively, appear to the right in equation (15). Thus, one can observe the substitution and output effects without matrix inversion and with greater computational facility.

Note that, if output is endogenous and expenditures are exogenous, there is no contradiction of underlying econometric assumptions when one derives the cost function from the indirect production function to obtain a substitution effect. However, if a cost function is initially estimated, assuming that output is exogenous, there is a contradiction if one attempts to do the reverse—that is, to derive the indirect production function from the cost function to calculate the output effect.

Consider the problem of output maximization subject to a cost or budget constraint with a two-input Cobb-Douglas production function (10). The first-order conditions may be solved algebraically to yield:

$$\begin{aligned}\bar{X}_1^* &= b_1 C P_1^{-1} \\ \bar{X}_2^* &= b_2 C P_2^{-1}\end{aligned}\quad (16)$$

Identical input demands may be obtained from the indirect production function $Q(P_X, C)$. For the indirect production function:

$$Q = C(b_0 P_1^{b_1} P_2^{b_2})^{-1} \quad (17)$$

and the input demands are:

$$\begin{aligned}X_1^* &= b_1 C P_1^{-1} = -(\partial Q / \partial P_1)(\partial Q / \partial C)^{-1} \\ X_2^* &= b_2 C P_2^{-1} = -(\partial Q / \partial P_2)(\partial Q / \partial C)^{-1}\end{aligned}$$

The Profit Function: A Dual Approach

Fuss and McFadden (13), Lau (21), and Normal and Russell (28) have shown that it is possible to postulate the existence of a profit function dual to a production function not characterized by constant returns to scale (for example, $F(X)$ in equation (10)), so that the locus of maximum profits for all P_Q and P may be written:

$$\Pi(P_Q, P) = \max_X P_Q F(X) - C(Q, P) \quad (18)$$

where $\Pi(P_Q, P)$ is the profit function. Profits become a function of output price and input prices. The profit function possesses the following properties:

- Positively linearly homogeneous in P_X and P_Q ,
- Quasi convex in P , and
- Decreasing in P , and increasing in P_Q .

For any production function $F(X)$ with these properties, a unique profit function exists. Similarly, given any profit function with the above properties, a unique production function can be shown to exist.

The salient feature of the profit function used in formulating the input demand function is that the optimal input demands can be obtained as the first derivatives of the profit function by use of Hotelling's lemma (18):

$$\frac{-\partial \Pi(P_Q, P)}{\partial P_i} = \hat{X}_i^*(P_Q, P) \quad (19a)$$

where $\hat{X}^*(P_Q, P)$ is the input demand function.

Moreover, by the same lemma, the supply function is given by:

$$\frac{\partial \Pi}{\partial P_Q} = Q(P_Q, P) \quad (19b)$$

where $Q(P_Q, P)$ is the product supply function. Using equation (19a), we may write:

$$\left. \frac{\partial \hat{X}_i^*}{\partial P_j} \right|_{dP_Q=0} = - \frac{\partial^2 \Pi(P_Q, P)}{\partial P_i \partial P_j} \quad (20)$$

where $(\partial \hat{X}_i^* / \partial P_j) |_{dP_Q=0}$ embodies the substitution, output, and maximizing effects of an input price change. The second derivative replaces the constrained optimization matrix inversion.

The separate effects of the input price change may be obtained by transformation similar to those used in transforming the indirect production function into a cost function to obtain substitution and output effects. The optimal profit level is defined as:

$$\Pi(P_Q, P) = P_Q Q(P_Q, P) - \sum P_i \hat{X}_i^*(P_Q, P) \quad (21)$$

where $Q(P_Q, P)$ is derived from the supply function (19b).

Solving implicitly for P_Q :

$$P_Q = H(Q, P) \quad (22)$$

Substituting this expression into the input demand equation, we then obtain:

$$\hat{X}_i^* = \hat{X}_i^*(H(Q, P), P) = \hat{X}_i^*(Q, P) \quad (23)$$

Equation (23) can be used to construct the cost function dual to the profit function at the profit-maximizing output level Q .

$$C = \sum P_i \hat{X}_i^*(Q, P) = C(Q, P) \quad (24)$$

Thus, the analyst can derive a cost function at output Q which can be used to obtain the substitution effect where $dQ = 0$.

$$\frac{\partial C(Q, P)}{\partial P_i \partial P_j} = K_{ij} \quad (25)$$

The analyst can also use $C(Q, P)$ to obtain the indirect production function $Q(C, P)$ and to obtain the output effect as in equation (15). The maximizing effect may be obtained by use of these transformations and equation (5c).

Thus, one can obtain the total effect and partial effects without inversion of a bordered Hessian matrix by exploiting the dual relationship among the profit, indirect production, and cost functions. Moreover, Mundlak's results can be used to decompose the effect for a partial equilibrium policy analysis. It is conceptually possible to obtain such effects for each variable input in production, with assurance of consistent aggregation. This process permits obtaining factor substitution and output adjustments from a well-defined profit function, supply function, and system of input demand functions.

Consider the problem of output maximization under a decreasing returns-to-scale Cobb-Douglas production function, subject to a fixed output price. The objective function becomes:

$$\text{Max}_X \Pi(P_Q, P) = P_Q F(X) - P_1 X_1 - P_1 X_1 - P_2 X_2 \quad (26)$$

Substituting equation (10) into (26), we derive:

$$\text{Max}_X \Pi(P_Q, P) = \frac{P_Q b_1^{-b_1} b_2^{-b_2}}{b_0} X_1^{b_1} X_2^{b_2} - P_1 X_1 - P_2 X_2 \quad (27)$$

where $b_1 + b_2 < 1$.

The profit function dual to $F(X)$ in (26) is written:

$$\Pi(P_Q, P) = b_0 P_Q^{b_1} P_1^{b_1} P_2^{b_2} \quad (28)$$

Solving equation (27) for $\hat{X}_1^*$ yields the same result as taking the derivative of the profit function with respect to input prices:

$$\hat{X}_1^* = -\frac{\partial \Pi(P_Q, P)}{\partial P_1} = -b_1 b_0 P_Q^{b_1} P_1^{b_1-1} P_2^{b_2} \quad (29)$$

The profit-maximizing input demands may be obtained through either a constrained optimization or a dual approach. However, the dual approach is analytically easier to implement when the production function is unknown and there are more than two inputs in the production process.

Functional Forms

Given the equivalence of the comparative statics effects and the nature of the dual technological characterizations, it is apparent that the analyst can proceed to derive input demand equations from cost, indirect production, or profit functions, depending on the assumed producer's objective function. The choice of specific functional forms for estimating and deriving input demand equations inevitably involves compromises. Fuss, McFadden, and Mundlak (14) list the following five criteria to help distinguish among forms:

1. Parsimony in parameters,
2. Ease of interpretation,
3. Computational ease,
4. Interpolative robustness within sample (concavity, monotonicity, and so forth), and
5. Extrapolative robustness outside sample (concavity, monotonicity, and so forth).

Among the more common functional forms providing a second-order approximation to any twice differentiable function are the following:

- Generalized Leontief (8),
- Translog (5),
- Generalized Cobb-Douglas (9),
- Quadratic (20),
- Generalized quadratic (28),
- Generalized concave (14), and
- Generalized Box-Cox (3).

The selection of a particular functional form for input demand analysis depends on the objectives of the analyst. These flexible functional forms often exhibit weakness in criteria 1, 3, and 5. The analyst must choose the form which best meets the objectives. For example, if the analyst wants to obtain input demand equations in a particular form such as input/output ratios, a generalized Leontief functional form may be more useful than the translog form. If technological augmentation is the primary objective, a translog may be superior. The generalized Box-Cox form serves as a parent function for both the generalized Leontief and the translog, but is exceedingly difficult to formulate for either the indirect production or profit functions. Nevertheless, it is most useful for testing maintained hypotheses concerning the degree of production

flexibility for cost functions. No one form serves all analytical purposes.

Applications of dual cost or profit functions intended to examine input demand include Norman and Russell (28), Berndt and Wood (4), Fuss and McFadden (13), and Christensen, Jorgenson, and Lau (5). Few applications of indirect production functions are available, but indirect utility functions are used by Houthakker (19) and by Christensen, Jorgenson, and Lau (5).

Conclusion

The use of cost, indirect production, and profit functions simplifies the discussion of changes in input demand because it implicitly embodies both the objective function and the production function. This approach requires few constraints on the form of the underlying technology from which the demand for inputs may be derived. These functions are particularly useful for estimating the effects of input price changes on input demand. Duality principles may be extended to problems of analyzing variable input demand with fixed inputs (28), technological change (3), uncertainty (29), and intertemporal decisions (2, 24).

Although the discussion of input demand is couched in a steady state equilibrium framework, Fuss and McFadden (13) and Berndt (2) have explored theoretical and empirical extensions of duality principles to evaluate production flexibility and disequilibrium considerations for analyzing input demand under input price changes. These approaches are particularly useful for many problems in which the production functions are unknown and steady state equilibrium assumptions may be questionable.

The duality approach still requires the researcher to make decisions concerning the variables under the control of the decisionmaker, the producer's objective function, and the appropriate characterization of technology from which input demand is derived. Although the duality approach provides a computationally more attractive alternative to input demand analysis, it requires similar assumptions concerning the nature of the technological characterization and the objective function as well as a comprehensive data base.

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Estimating Input Cost Shares for Agriculture Using a Multinomial Logit Framework

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Abstract

Many econometric analyses include dependent variables constrained to the interval between zero and 1. Under such conditions, simple regression procedures break down. Several alternative stochastic models which avoid this problem can be defined depending on the assumed error structure. Two alternative forms of the logit model are treated here. The multivariate logit approach assumes that the share specification is an accurate representation of the underlying input demand structure. The multinomial logit approach treats the dependent variable as a probability with a multinomial density.

Keywords

Econometrics, limited dependent variable, multinomial logit, maximum likelihood

Many econometric analyses include dependent variables which are constrained to the interval between zero and 1. Typical of this type of analysis is the simultaneous estimation of input cost shares. Development of the translogarithmic (translog) cost function has increased interest in estimating systems of input share equations (10, 11). Cost functions and underlying share equation systems have been estimated by Christensen and Green (9), Berndt and Wood (6), Denny and Pinto (12), and Humphrey (18).¹ From share equations it is possible to derive input price and substitution elasticities (5, 20, 25).

However, there is no implicit or explicit mechanism constraining the prediction of input shares to between zero and 1 by use of simple regression procedures. Predictions may fall outside the zero-1 interval and, because of the grouped nature of the data, error terms are likely to be heteroscedastic. The objective here is to outline two versions of a logit model which explicitly force predicted input shares to sum to 1. The logit is a sensible and convenient alternative to the limited dependent variable problem encountered when one estimates input share equations derived from a translog cost function.

First, I describe the underlying structure of the logit model. I present two alternative forms of the logit model and discuss an estimation methodology for each. Finally, I cite an example of a multinomial logit model where a maximum likelihood technique is used to estimate input cost shares for agriculture.

The Logit Model

The logit model is one approach to the analysis of a general class of problems termed discrete choice behavior. These types of problems have been investigated in other fields for many years (2, 3). Economists have begun to analyze problems which could be cast in this framework by considering discrete choices as selections from a continuum of alternatives, thereby integrating these problems into the theory of the household and firm.

Statistical economic analysis of the general population or the choice behavior of the average firm or consumer is complicated because such behavior must be described in probabilistic terms. The probability of a particular choice is conditional on the explanatory variables selected for the analysis. A discrete choice framework has been used to examine many problems including migration (28), occupational choice (7, 30), demand for housing (35), and demand for consumer durables (16, 33).

Interest in the logit derives from its relatively simple structural form (29). The logit forces estimates of the dependent variables to be between zero and 1 while summing to unity. It allows for a wide array of functional forms, although certain functions are more easily applied. When applied in consumption analyses, the logit allows for nonunitary income elasticities, and when applied in production analyses, it does not place *a priori* restrictions on elasticities of substitution. Recent applications of the multinomial logit analyze the expenditures of firms on inputs (20) and of consumers on household goods (34).

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¹ Italicized numbers in parentheses refer to items in the References at the end of this article.

The basic structure of the logit expenditure system is written as:

$$S_i = \exp(c_i(x)) / \sum_{j=1}^k \exp(c_j(x)) \quad (i=1, 2, \dots, k) \quad (1)$$

where S_i is the share or proportion of total expenditures spent on the i th good, x is a vector of explanatory variables such as prices or income, and c_i is any mathematically well-behaved function linking the explanatory variables and the dependent share variables.

Two alternative forms of the logit transformation can be used for empirical analysis. The multinomial form of logit, or conditional logit (27), treats the dependent variable as a probability with a multinomial density. The multivariate logit emerges from a more *ad hoc* prediction-oriented approach which assumes that equation (1) is an accurate functional characterization of expenditures. With either form, one can incorporate many factors which influence expenditures, such as family or firm size, as explanatory variables while maintaining a theoretically consistent expenditure system.

Multinomial Logit

The theoretical foundation of the multinomial logit sets the decisionmaker in an environment where discrete choices of expenditures are made. If it is assumed that the larger the value of an index the greater the probability that the event will occur, then one can define a monotonic relationship linking the value of the index and the probability of the event's occurrence.

The decisionmaker's choices are described in a decision index of the form:

$$c = c(x) + e \quad (2)$$

where e is a random disturbance associated with a given probability distribution and $c(x)$ is nonstochastic. Faced with k alternatives, the decisionmaker will choose alternative i only if $c_i(x) + e_i > c_j(x) + e_j$ for all $j \neq i$. The probability of this event occurring is:

$$P_i = \text{Prob}[c_i(x) + e_i > c_j(x) + e_j] \quad \text{for all } j \neq i \quad (3)$$

and, therefore:

$$P_i = \text{Prob}[e_i - e_j > c_j(x) - c_i(x)] \quad \text{for all } j \neq i \quad (4)$$

One must make binary comparisons of $c_i(x)$ with each alternative $c_j(x)$. Probability distributions which are closed under subtraction, or produce convenient distributions when subtracted, are particularly attractive candidates for the probability densities of e (19).

If $F(e_1, e_2, \dots, e_k)$ represents the cumulative distribution function of the disturbances e_i and if F_i denotes the marginal density function of e_i (the derivative of F with respect to e_i), then the probability P_i is written:

$$P_i = \int_{-\infty}^{\infty} F_i[e_1 + c_1(x) - c_i(x), \dots, e_k + c_k(x) - c_i(x)] de_i \quad (5)$$

If the errors are independent and identically distributed, then a necessary and sufficient condition for the model described by equation (4) to yield the conditional logit form is that the errors have a Weibull density (27). That is:

$$e_i \sim \exp(-\exp(-e_i)) \quad (6)$$

The Weibull density is a convenient way to generate the logistic distribution. It is closed under subtraction and closely approximates the normal distribution while being numerically simpler. McFadden (27) indicates that the underlying choice structure implies the independence of irrelevant alternatives axiom. The independence of irrelevant alternatives condition is both a strength and a weakness of the logit expenditure model (14). It is a strength because introducing additional alternatives does not alter the relative odds with which previous alternatives are selected. It is a weakness because it requires that the cross-elasticity of demand for each old expenditure category, with respect to an attribute of a new category, is uniform across all old categories.

If the multinomial form of the logit is used to describe producer expenditures on inputs, then a set of k independent conditional probabilities are assumed to jointly determine the allocation of expenditures into k input categories. The probabilities have a logistic structure and are conditional on input prices and other explanatory variables. The probabilities are written as:

$$P_i = \exp(c_i(x)) / \sum_{j=1}^k \exp(c_j(x)) \quad (i = 1, 2, \dots, k) \quad (7)$$

where P_i is the conditional probability of \$1 being allocated to input i , x is a vector of explanatory variables, and $c_i(x)$ is a decision index.

Because the probability P_i is unobservable, the model is made operational by use of the share of the cost of production associated with input i as a proxy (31). Equation (7) is therefore rewritten as:

$$S_i = v_i q_i / M = \exp(c_i(x)) / \sum_{j=1}^k \exp(c_j(x)) \quad (i = 1, 2, \dots, k) \quad (8)$$

where M is the sum of each input (q_i) used in production multiplied by its input price (v_i).

Before the multinomial logit input expenditure model can be estimated, a functional form must be selected for the decision index $c_i(x)$. The function may include input prices, output, and all other production or producer characteristics considered relevant. Although no specific restrictions on $c_i(x)$ are necessary to estimate the multinomial logit model, estimation is easier if the functions are assumed to be linear in their parameters and invariant in structure between equations except for interequation parameter variability. The functions are written as:

$$c_i(x) = \sum_{j=1}^n B_{ij} h_{ij}(x) \quad (i = 1, 2, \dots, k) \quad (9)$$

where B_{ij} are parameters and h_{ij} are functions.

The expenditure of \$1 on a given input is analogous to sampling with replacement from a population classified into k categories. The resulting multinomial distribution is written as:

$$f = \frac{M!}{E_1! \dots E_k!} P_1^{E_1} \dots P_k^{E_k} \quad (10)$$

where E_i equals $v_i q_i$ and is non-negative. For a sample of T observations, the logarithm of the likelihood function associated with equation (8) is written as:

$$\ln L = \text{Constant} + \sum_{t=1}^T \left(\sum_{i=1}^k S_{it} c_{it} - \ln \left(\sum_{i=1}^k \exp(c_{it}) \right) \right) \quad (11)$$

Although the functions $c_i(x)$ may take any form, functions which are linear in parameters are sufficiently flexible for most purposes. A linear form also leads to simple expressions for the maximization of the likelihood function. Therefore, if:

$$c_{it} = \sum_{j=1}^n B_{ij} x_{jt} \quad (i = 1, 2, \dots, k) \quad (12)$$

then, the first-order conditions for the maximization of the log-likelihood function are:

$$\frac{\partial \ln L}{\partial B_{ij}} = \sum_{t=1}^T x_{jt} [S_{it} - P_{it}] = 0 \quad (i = 1, 2, \dots, k) \quad (13)$$

$$(j = 1, 2, \dots, n)$$

The second-order conditions for the maximization of the log-likelihood function are:

$$\frac{\partial^2 \ln L}{\partial B_{ij} \partial B_{ij'}} = \sum_{t=1}^T x_{jt} x_{j't} P_{it} (P_{it} - 1) < 0 \quad (i = 1, 2, \dots, k)$$

$$(j = 1, 2, \dots, n) \quad (14)$$

where both parameters (denoted by j and j') are in the i th equation and:

$$\frac{\partial^2 \ln L}{\partial B_{ij} \partial B_{i'j'}} = \sum_{t=1}^T x_{jt} x_{j't} P_{i't} P_{i't} < 0 \quad (i = 1, 2, \dots, b)$$

$$(j = 1, 2, \dots, n) \quad (15)$$

where each parameter is from a different equation.

The maximum likelihood estimators are invariant to monotonic transformations of the likelihood function. The log-likelihood function is maximized where the first-order derivatives are zero (equation (13)) and the matrix of second-order derivatives, formed by the derivatives in equations (14) and (15), is negative definite. Because the first and second derivatives are nonlinear in parameters, an iterative search procedure is needed to solve for the parameters B_{ij} .

The estimation procedure is complicated by an indeterminacy in the equation system which arises because the sum of the shares must equal 1. The share equations are, therefore, invariant with the addition of the same expression, $\ln Z$ for example, to each decision index. The indeterminacy causes the matrix of second-order partial derivatives of the log-likelihood function to be singular.

One can avoid the singularity problem by normalizing the parameters for a particular variable in the k functions (34). The normalization does not affect the predicted shares. A straightforward approach is to divide the $k-1$ equations by the k th equation. In a logarithmic form, the share system is:

$$\ln(S_i/S_k) = \sum_{j=1}^n (B_{ij} - B_{kj}) h_{ij}(x_j) \quad (i = 1, 2, \dots, k-1) \quad (16)$$

Maximizing $\ln L$ with respect to c_i is equivalent to maximizing $\ln L$ with respect to $c_i - c_k$ for any k (34). It is only necessary to consider $k-1$ equations as the information provided by the k th equation is constant. Furthermore, regardless of the normalization employed, the predicted values of S_i are identical. Therefore, equation (16) can be used to estimate input cost shares by use of linear regression. One can compute the individual input cost shares, S_{ij} , from the regression results while forcing the shares to sum to unity.

Multivariate Logit

An alternative stochastic form of the logit uses the share system given in equation (8) directly. The share equations are interpreted as an accurate characterization of producer expenditures for inputs (8, 22). One can generate the stochastic model structure by appending error terms e_i to each share equation.

Each equation's error term represents deviations between optimal cost shares and observed cost-minimizing shares. There are several reasons for the existence of the error term: the failure of input markets to clear perfectly, the aggregation or measurement error, or the randomness of human behavior. The error term associated with the i th cost share is assumed to be distributed normally with mean zero and variance σ_i . The variance is not assumed to be constant across shares because the variance of e_i is generally not equal to the variance of e_j .

Three covariances are generated by the error terms in the share equations. One of these is the covariance between disturbances of different observations and of the same share equation:

$$\sigma_{ii} = E(e_{it}, e_{it'}) \quad (t \neq t' = 1, 2, \dots, T) \quad (i = 1, 2, \dots, k) \quad (17)$$

The second represents the relationship between different share equations and observations:

$$\sigma_{ij} = E(e_{it}, e_{jt'}) \quad (t \neq t' = 1, 2, \dots, T) \quad (i \neq j = 1, 2, \dots, k) \quad (18)$$

Both these covariances are assumed to equal zero.

The third covariance arises from the combination of different share equations and the same observation:

$$\sigma_{ij} = E(e_{it}, e_{jt}) \quad (t = 1, 2, \dots, T) \quad (i \neq j = 1, 2, \dots, k) \quad (19)$$

This covariance is usually referred to as the contemporaneous covariance (32). Because the underlying production structure is estimated as a system, it is unlikely that the contemporaneous covariances are zero. By appropriately stacking the share equations, one can write the variance-covariance matrix for the disturbance term as a block-diagonal matrix with T diagonal submatrices.

This matrix represents the interdependency of the k share equations for each observation ($t = 1, 2, \dots, T$). The off-diagonal submatrices of the error system's variance-covariance matrix are zero by assumption.

The specified system of share equations is characterized as a seemingly unrelated regression problem (41). Zellner's generalized least-squares procedure cannot be directly applied, however. Because the disturbances of the share equation must sum to zero, the estimated variance-covariance matrix nec-

essary for implementing Zellner's procedure is singular. One can transform the share equations to an estimable form by normalizing with the k th share equation.

Parameters estimated by the Zellner generalized least-squares procedure are not invariant to the choice of common denominator share when an estimated variance-covariance matrix is employed. However, maximum likelihood estimates are invariant to which equation is deleted (1), and iterating the Zellner procedure leads to maximum likelihood estimates (13, 21). Therefore, iterating the Zellner procedure is a computationally efficient means of deriving parameter estimates. In general, the properties of maximum likelihood estimators only hold asymptotically. However, most of the maximum likelihood estimators' asymptotic properties are present in small samples (21).

If the same set of regressors is utilized in all $k - 1$ equations, then the iterative Zellner procedure and ordinary least squares give identical parameter estimates. However, estimates of the standard errors may differ.

An Example of a Multinomial Logit Estimation

A multinomial logit model is used to estimate a system of cost share equations for agricultural inputs. In addition, the price elasticity of demand for each input is calculated by use of the multinomial logit parameter estimates. Although the procedure allows one considerable flexibility in selecting a functional form and avoids the limited dependent variable problem, it is still subject to the same practical difficulties (for example, data aggregation) in applying all econometric models.

Estimated Form

The underlying model structure estimated in this example is:

$$S_i = \exp(a_i + \sum_{j=1}^k B_{ij} \ln v_j) / \sum_{i=1}^k \exp(a_i + \sum_{j=1}^k B_{ij} \ln v_j) \quad (20)$$

where S_i is the i th input share, a and B are unknown parameters, and v_j is the price of input j .

Because the model specified in equation (20) leads to a singular matrix of second-order derivatives, a maximum likelihood procedure cannot be applied. Instead, one can transform the share system to an estimable form by normalizing on the k th equation. After taking logarithms, the estimated form of the share system is:

$$\ln(S_i/S_k) = (a_i - a_k) + \sum_{j=1}^k (B_{ij} - B_{kj}) \ln v_j \quad (i = 1, 2, \dots, k - 1) \quad (21)$$

The parameters a_i and B_i are now defined as differences from the parameters of the k th equation.

Data and Estimating Methodology

Production cost data were developed for land, labor, fertilizer, energy, and capital inputs for the United States. Cross-sectional data for 39 continental States for 1974 are utilized (15, 24, 38, 39, 40).

Because the likelihood function associated with equation (21) has non-linear first and second derivatives, an iterative search procedure is used to solve for the maximum likelihood estimates. Tyrrell (34) developed the computer software used in this analysis to analyze household expenditure patterns. Tyrrell employs both first and second derivatives in an extended Newton-Raphson procedure (see (17, 26)). The final output includes a vector of estimated coefficients, first- and second-order derivatives of the likelihood function, asymptotic standard errors, and asymptotic t-statistics.

Model Results

I estimated the expenditure system by normalizing on the capital cost share. Table 1 shows the estimated parameters and asymptotic standard errors of the model specified in equation (21). Because maximum likelihood estimates are asymptotically normally distributed and the standard errors are asymptotically distributed as a chi-square, the ratios of the estimated parameters to their standard errors can be interpreted as asymptotic t-statistics.

The t-statistics indicate general support for the estimated parameters. Of the 24 parameter estimates, 16 exceed 1.0; of these, 10 exceed 2.0. The t-statistics for the own-price parameters for each input are the highest. The land and capital price parameter for each input are the lowest. The low-price land and capital parameters may be associated with poor measurement of these inputs.

Because the share system is indeterminant, it is not possible to identify each B_{ij} in the expenditure system. The estimated

parameters can only be interpreted as differences from the parameter associated with the normalizing input cost share (see equation (23)).

The estimated likelihood function can be used to formulate statistical tests (4). The test statistic, $-2\ln\lambda$, is asymptotically distributed as a chi-square with degrees of freedom equal to the number of independently imposed restrictions when:

$$\lambda = (|\Omega_R|/|\Omega_U|)^{-T/2} \quad (22)$$

where $|\Omega_R|$ and $|\Omega_U|$ are the determinants of the restricted and unrestricted estimated variance-covariance matrices of error terms (32). The chi-square test allows for comparisons of different models if the restricted model is nested within an unrestricted model. In this analysis, the chi-square test is used to test the null hypothesis that all coefficients equal zero. The null hypothesis is easily rejected at the 1-percent level with 30 degrees of freedom.

Although the estimated share system is normalized on the k th share equation (capital), it is possible to solve for the predicted shares for all inputs. The k th share can be computed because:

$$1/(\hat{S}_k + \sum_{i=1}^{k-1} \exp(\ln(\hat{S}_i/\hat{S}_k)\hat{S}_k)) = 1 \quad (23)$$

where $\hat{S}_k$ is the projected cost share for the k th input (capital) and $\ln(\hat{S}_i/\hat{S}_k)$ is the estimated dependent variable associated with the i th equation. The estimated share for the k th input is computed as:

$$\hat{S}_k = 1/(1 + \sum_{i=1}^{k-1} \exp(\ln(\hat{S}_i/\hat{S}_k))) \quad (24)$$

Table 2 reports the predicted and observed cost shares. The model fails to predict actual cost shares with a high degree of accuracy. The average absolute differences between predicted and observed input cost shares are 0.05 (land), 0.07 (labor),

Table 1—Multinomial logit parameter estimates¹

Input	Input price					
	Land	Labor	Fertilizer	Energy	Capital	Intercept
Land	0.816 (.117)	0.199 (.148)	0.111 (.033)	-0.390 (.280)	-0.653 (.726)	-7.420 (2.450)
Labor	-.009 (.141)	1.233 (.178)	-1.320 (.358)	1.384 (.346)	.632 (.827)	-1.040 (2.842)
Fertilizer	.121 (.129)	.274 (.166)	.576 (.359)	.806 (.322)	.031 (.785)	-8.330 (2.700)
Energy	-.023 (.183)	.987 (.242)	1.249 (.492)	-.496 (.370)	-.569 (1.092)	-6.350 (3.546)

¹ Standard errors in parentheses.

0.05 (fertilizer), 0.02 (energy), and 0.03 (capital), whereas the average observed cost shares are 0.28 (land), 0.17 (labor), 0.20 (fertilizer), 0.08 (energy), and 0.28 (capital). If the absolute differences are compared with the observed shares, then labor and energy are the least accurate predictions and capital is the most accurate. The bad predictive capability of the model is largely attributed to the use of aggregate cross-sectional data. The appropriate unit of observation is the firm. However, no reliable set of input price and quantity data are available at the firm level. This problem is particularly true for energy.

In addition to predicting the individual cost shares, the multinomial logit model enables the analyst to derive the price elasticities implicit in the derived demand for each input. If it is assumed that total expenditure on inputs, M , is not invariant to changes in input prices, then the own-price and cross-price elasticities are approximated by:

$$E_{ii} = S_i - v_i(\partial f_i / \partial v_i - \sum_{j=1}^k S_j \partial f_j / \partial v_i) - 1$$

$$(i = 1, 2, \dots, k) \quad (25)$$

and by:

$$E_{ij} = S_i - v_j(\partial f_i / \partial v_j - \sum_{j=1}^k S_j \partial f_j / \partial v_j)$$

$$(i \neq j = 1, 2, \dots, k) \quad (26)$$

The indeterminacy caused by the adding-up constraint prevents direct calculation of the input price elasticities. However, the predicted price elasticities can be derived if the elasticities associated with the k th share are derived first. The predicted elasticity for the k th share with respect to the i th price is:

$$\hat{E}_{ki} = \hat{S}_k - \sum_{j=1}^{k-1} \hat{S}_j (B_{ji} - B_{ki}) + -1 \text{ (if } i = k; \text{ or } 0 \text{ if } i \neq k)$$

$$(i = 1, 2, \dots, k - 1) \quad (27)$$

and, therefore:

$$\hat{E}_{ji} = \hat{E}_{ki}(B_{ji} - B_{ki}) \quad (i \neq j = 1, 2, \dots, k - 1) \quad (28)$$

Table 3 reports the average predicted own-price and cross-price elasticities. All the own-price elasticities have the right sign except labor, which is close to zero. The magnitudes of the own-price elasticities are similar to a translog specification with the same data (24). They do, however, differ significantly from translog specifications estimated with time-

Table 3—Average price elasticities of predicted input demand

Factor	Input price				
	Land	Labor	Fertilizer	Energy	Capital
Land	-0.152	0.116	1.051	-0.559	-0.623
Labor	-.120	.006	-1.524	1.072	.519
Fertilizer	.099	.135	-.539	.583	.008
Energy	.027	.921	1.206	-1.647	-.521
Capital	-.022	-.227	-.115	-.152	-.600

series data (23). This discrepancy suggests that the data and not the model specification are the source of the elasticity differences.

Unlike a translog system, cross-price elasticities are not constrained to be symmetric. In fact, the cross-price elasticities for land and labor and for land and energy do not have the same signs. The cross-price elasticities indicate that land is a substitute for labor and fertilizer in farm production, but is a complement with energy and capital (36). Controversy still exists over the relationship of energy to other inputs in manufacturing. The results here indicate that energy is a complement with capital and is a substitute with labor in agricultural production. Although other analyses confirm the results for capital, results for labor are ambiguous and depend on whether the model distinguishes between long- and shortrun input substitution (23).

Summary

The logit model provides a flexible alternative to the more popular translog approach to estimating systems of input share equations. The logit allows for a wider range of explanatory variables and functional forms than does the translog. Furthermore, the logit constrains share predictions to the interval between zero and 1. Two forms of the logit (multivariate and multinomial) can be defined depending on the assumed error structure. For input expenditure systems, the multivariate logit assumes the share specification accurately represents the underlying input demand structure. The multinomial logit treats the dependent variable as a probability with a multinomial density. Either model can be estimated with well-known techniques and can provide estimates of predicted input cost shares and price elasticities of demand. However, because of the indeterminacy of the share system, individual parameter estimates can not be identified for the multinomial logit model.

The logit model has been applied to many economic problems. Within the context of agricultural production, additional applications might include dynamic models or incorporation of the cost share system into a macromodel where agricultural production is only one of many production sectors. Because the logit's structure forces share systems to sum to unity, it is an attractive candidate for maintaining consistent predictions.

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Research Review

Tax Burdens in American Agriculture: An Intersectoral Comparison

Charles Adair Sisson. Ames: Iowa State University Press, 1982, 154 pp., \$14.40.

Reviewed by Wendy Rome, James Hrubovcak, and Ron Durst*

Major changes in the agricultural sector have prompted many observers to question the necessity and desirability of maintaining specific tax provisions which benefit farmers—for example, cash accounting; expensing capital costs for groves, orchards, and vineyards; and allowing farmers to treat gains on sales of qualifying livestock held for draft, breeding, dairy, or sporting purposes as capital gains. Many economists and tax experts believe that farmers no longer need the preferential treatment contained in the Internal Revenue Code. Others claim that farming has not changed substantially and that farmers still need these provisions. Still others suggest that changes in the agricultural sector and higher land values have created new problems for farmers that warrant additional tax relief.

Charles Sisson's book investigates comparative farm and non-farm burdens. It outlines for tax researchers and policymakers the fundamental issues of farm taxation. Sisson discusses some important issues of farm taxation and states several hypotheses. First, he suggests that the tax burden for the farm population is substantially lower than it is for the U.S. population as a whole. Second, he hypothesizes that this viewpoint is based on the fact that farmers receive preferential treatment under Federal income tax laws. Third, he suggests that the difference between farm and nonfarm tax burdens is attributable to the extremely low tax burdens of high-income farmers. Finally, he questions whether Federal estate taxes impose a greater tax burden on farmers than on the total population so that these taxes should be considered in the overall analysis of comparative farm and nonfarm tax burdens.

Sisson resolves definitional, methodological, and data problems inherent in farm tax research. He presents his results and evaluates whether his analysis is valid without recognition of the effects of the estate tax. Finally, he recommends areas for tax reform.

Potential tax liabilities can significantly affect investment decisions. Sisson demonstrates that preferentially taxed farm investments have higher after-tax present values than other investments. Thus, tax preferences may significantly affect investment decisions and possibly lead to distortions in the allocation of farm resources. Have investors shifted investment in an effort to maximize their after-tax incomes? To what extent are tax preferences in the farm sector encouraging nonfarm investors to enter agriculture? Sisson examines

the issue of tax induced capital in agriculture and attempts to determine the significance of tax-loss farming.

Two fundamental questions must be resolved before we examine the farm-nonfarm tax burden issue. First, what is a farmer? Second, how do we measure tax burdens? "Farmer" is an inherently ambiguous concept. Citing the changing nature of the farm sector and the need for a set of farm definitions that policymakers can use for multiple purposes, Sisson concludes that a single definition is no longer sufficient to identify the farm population. Sisson uses five definitions of a farmer that are based on economic and socioeconomic criteria. These include the Bureau of the Census and the Internal Revenue Service (IRS) definitions as well as those based on income and place of residence. Examining the farm-nonfarm tax burden issue for various farm definitions provides valuable information about the farm population and insight into the probable effects of tax reform.

The method by which tax burdens should be measured is a somewhat more difficult issue. Sisson concludes that the only feasible alternative is the ability-to-pay approach measured by the Haig-Simons concept of income, a concept which he fully develops in the book. Sisson uses a simple tax burden index consisting of tax liability divided by Haig-Simons income. The tax liability portion of that index was determined on the basis of apparent tax liabilities and of alternative progressive and regressive tax incidence assumptions. Sisson's progressive scenario generally assumes a tax is borne by property income whereas his regressive alternative assumes that consumers bear the tax.

Sisson's results depend on the data set and his assumptions. We have already mentioned many of his assumptions, but have not discussed his use of the Brookings MERGE synthetic data set. The MERGE file is a synthetic data set compiled from cross-matching 87,000 records from the IRS tax file for 1966 with 30,000 household records from the 1967 Survey of Economic Opportunity (SEO). Ideally, combining these two data sets would provide a close approximation of the tax returns filed in calendar year 1966 by each of the 30,000 households in the SEO data file. If matched properly, the more complete tax information provided by the IRS data combined with the demographic information included in the SEO data file would be a more powerful data base than either data file used alone. This file was described in detail in an earlier issue of this journal (1).¹ In the MERGE file, less than 2 percent of the SEO data file was not matched with IRS

*The reviewers are agricultural economists with the National Economics Division, ERS.

¹ Italicized numbers in parentheses refer to items in the References at the end of this review.

data, and these records were generally concentrated in the highest income classes. Sisson evaluated his results, excluding the unmatched data in the demographic data set (limited to families with less than \$80,000 in adjusted gross income) and including the unmatched data in the income data set. Sisson believed that the comparison of the results from the demographic and income data sets could provide information on the tax treatment of various components of the farm population and on their distribution.

Sisson compares the tax burden by adjusted family income of the total population to the tax burdens under the five alternative definitions of a farmer and under varying progressive and regressive tax-incidence assumptions. Except for the lowest income classes, farm families have a significantly lower tax burden than the total population irrespective of the definition of farm or the incidence assumption. Sisson concludes that the benefits of preferential tax treatment for farmers are fairly evenly distributed across the rural sector.

Sisson compares GINI ratios based on adjusted family income before taxes to measure the inequality of the income distribution within the farm sector in comparison with the total population.² He concludes that the income distribution is much more evenly distributed in the total population. He next calculates GINI ratios based on after-tax adjusted family income and concludes that, because the GINI ratios fall for both the total population and the alternative definitions of a farmer, the tax system does improve distributional equality. Sisson also calculates GINI ratios based on after-tax income with the capital gains exclusion eliminated. He concludes that, as the GINI ratios are not significantly affected, the repeal of the capital gains provisions would do little to improve distributional equality. The benefits of the capital gains provisions are broadly used across all income classes within the farm sector.

Sisson addresses the question of tax-loss farming by analyzing the distribution of farm losses of those individuals whom he considers more dependent on farming as a primary occupation than others who are only marginally involved farming. Although he acknowledges the definitional problems encountered in dividing these subgroups, he nonetheless makes two important conclusions. First, the great bulk of farm losses are incurred by individuals who depend heavily on farming for income. Second, "while tax-loss farming is not pervasive in agriculture, neither is it rare." Therefore, he concludes that tax preferences for the agricultural sector provide a strong incentive for high-income taxpayers to invest in agriculture.

Sisson attempts to determine whether the effects of the estate tax on farmers and nonfarmers warrant any changes. If estate

tax burdens in 1966 varied significantly between farm estates and nonfarm estates, then the comparative tax burden analysis would have to be modified. Until 1977, all estates faced the same estate tax laws. In 1976, however, the Congress enacted two special provisions for estate taxes that benefit farmers. Farmers used two basic arguments in 1976 to convince the Congress that they needed preferential estate tax treatment. First, they argued that the estate tax was based largely on land values which were unrealistically high relative to the land's income-generating ability. Second, they argued that farm liquidity was lower than that in other types of estates and that the estate tax was, therefore, hard to pay.

Sisson suggests that if farmers needed this special assistance in 1976, then perhaps they needed it in 1966, the year on which he based his analysis. He attempts to determine whether or not the estate tax burden for farmers was higher than it was for nonfarmers in 1966.

Using the MERGE file for 1966, Sisson compares liquidity-to-estate ratios to determine relative farm-nonfarm estate tax burdens. Whereas low liquidity was certainly a fundamental argument supporting the enactment of preferential estate tax provisions for the farm sector, it is not the appropriate measure for determining relative estate tax burdens between farm and nonfarm sectors. Rather, Sisson seems to be measuring a ratio of "undue hardship" as manifested in insufficient or barely sufficient liquid assets available to pay the estate tax at the time the tax is imposed.

Sisson concludes that the relative estate tax burden in 1966 in the farm sector vis-a-vis the total population was not sufficiently different to warrant adjusting his original analysis to include the estate tax. Although his findings do not actually support this assertion, because of his failure to measure actual estate tax burdens, this conclusion may, nevertheless, have been true in 1966.

In his final chapter, Sisson discusses the limitations of his analysis. At least two of these are significant enough to cast doubt on his conclusions. Our most serious concern is his failure to differentiate between personal income earned as wages or professional fees and personal income earned from the ownership of a business. The Internal Revenue Code is saturated with tax provisions designed to stimulate business investment and activities. Such provisions are generally not available to wage earners. Thus, farmers and other business owners would be expected to incur lower tax burdens than the total population, especially in the Federal tax area to which Sisson attributes a significant portion of the tax burden differential. Sisson does not compare the tax burden of farmers with the burden of other business owners.

Another limitation is the relative inability of the MERGE data file to depict current economic conditions. Sisson concludes that "the observations regarding farm-nonfarm tax

² The GINI coefficient is an index used to measure inequality in the distribution of income. The GINI coefficient ranges from zero (perfect equality) to 1 (absolute inequality).

burdens are nearly as valid today as they were for 1966.” He cites a statement by Joseph Pechman in 1975 regarding the continued validity of the 1966 MERGE file as support. If Sisson means the year 1975 when he uses the word “today,” then his conclusion is less subject to question. However, if he is speaking of the early eighties, serious doubts arise both because of extensive Federal, State, and local tax reform and because of changing economic conditions in the farm sector. We, therefore, must agree with the author’s statement that his tax differential estimates “must be regarded as upper bounds for the effects of these special tax rules rather than unbiased estimates of their amounts.”

Despite the limitations of Sisson’s study, it is useful to tax researchers primarily because of its contributions to the use

of synthetic microdata files in tax research and its discussions on fundamental farm tax research issues. For researchers less familiar with farm tax issues but who have some interest in these issues, the book provides substantial insight into those areas which, in the past decade, have dominated research on farm taxes.

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Asian Village Economy at the Crossroads: An Economic Approach to Institutional Change

Yujiro Hayami and Masao Kikuchi. Baltimore:
The Johns Hopkins University Press, 1982, 275 pp., \$25.00.

Reviewed by Ray W. Nightingale*

The introduction of modern cereal varieties (MVs) into rural Asian communities with rapid population growth produces economic and social change. The authors examine the thesis that MVs cause economic polarization; they also examine Ruttan's thesis that altered resource availabilities induce institutional change (1).¹ The authors ultimately absolve MVs of blame for polarization. They find that changes in returns to resources are brought about by institutional change induced by failures in factor markets. Although the evidence for this conclusion is drawn principally from four village case studies, they argue that these findings hold widely in Asia (as the title suggests).

To draw an Asian conclusion from four village case studies, Hayami and Kikuchi take an advocate's approach. They thoroughly acquaint themselves not only with these villages but also with the body of evidence on rural institutional change in Asia. They informally, but rigorously, build their case. They do this in a thoroughly professional manner, treating the reader to a highly rewarding exploration of critical agricultural development issues and the down-to-the-village facts.

In chapter 2, Hayami and Kikuchi describe their economic approach to examining institutional change, defined as "an attempt to analyze, by the logic of economics, the process by which institutional change was dictated by noneconomic as well as economic forces." Economic determinism is specifically disavowed. They argue that the distribution of income among various resource contributors (landlords, tenants, and laborers) responds to changes in resource endowments and technology through adjustments in institutions governing the use of land and labor. The following are some examples of the application of economic logic to these problems:

A person is altruistic to the extent that the return to altruism exceeds the cost of behaving as an altruist. . . .villagers will violate the village institutions if gains exceed costs.

If the [enterprising village leader] senses that the outcome will be more efficient if each member of the group pays a share of the marginal cost of additional units of the collective good equal to his share of the benefits. . .and others do not sense this, the leader will be able

to suggest arrangements which can leave everyone in the group better off.

In chapter 3, Hayami and Kikuchi give a macro and micro overview of agrarian change in Asia. They present the economics of income distribution among laborers and asset holders under increasing population and changing technology. They map two Asian roads of change: (1) polarization of peasant communities into commercial farmers and landless wage-workers, and (2) stratification of communities into a continuous spectrum of social arrangements for organizing labor and assets—from landless laborers and noncultivating landlords, to sharecropper, to peasant proprietors. With polarization, traditional patron-client mechanisms are replaced by impersonal market relations. With stratification, the personal mechanisms are sustained. The basic principle supporting the author's economic approach is that scarce resources result in a tight community structure which fixes economic roles, whereas informal markets require informal social structures.

The remainder of the text (200 pages) is devoted to examining whether polarization or stratification predominates and, on finding that stratification does, to exploring whether peasant polarization can be forestalled under increasing population pressure. The conclusion is that polarization can be avoided if efforts to generate scale-neutral technological progress (such as MVs) continue.

The polarization/stratification conclusions are based on a review of evolving agrarian institutions in Luzon (the Philippines), a survey of rice harvesting methods in 100 villages, and two village case studies. The authors base their conclusion on prospects for continued adjustment to population pressure through peasant stratification on a review of agrarian change in Java, a survey of rice harvesting methods in 48 villages, and two village case studies (one technologically stagnant and the other technologically progressive).

This book is a major step forward. It brings the insights of various disciplines to bear on the forces governing village change. The authors, both agricultural economists, enter the terrain of other disciplines and extend their own precepts into new realms. There is always a risk that in doing so one will lose or misrepresent the contribution of others, building up rather than tearing down barriers to understanding. It is clear that they took great care to avoid this, with solid success. They demonstrate a wealth of understanding of village society and institutions, and they clearly demonstrate how provision is made for resource organization and management.

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¹ Italicized numbers in parentheses refer to items in the References at the end of this review.

Through their research approach and their demonstrated depth of understanding, the authors have broken new ground and have produced a thoroughly readable book. It is enlivened by a common theme that runs throughout: Marxian prophecy of the violent course of human events as population presses on land and its nonfulfillment in this case. According to Marx, "At a certain stage of development, the material forces of production in society come into conflict with the property relations within which they had been at work before." The authors share with Marx an awareness of this conflict, but not his notion of an abrupt and complete transformation of the entire institutional framework. Village stratification is the revealed mechanism for resolving that conflict. Radical political economists are identified as perpetrators of the notion that new technologies (such as MVs) are major factors in promoting inequality and class differentiation.

I find the authors' verdict of the MV's absolute innocence of this charge welcome news indeed; however, they might have avoided highlighting the implications for radical economic doctrine and prophecy. Doctrines long held succumb only to overwhelming evidence to the contrary. The authors' building-block approach is the right one at

this stage for advancing understanding of a complex and challenging problem. But their approach is inadequate to the task of challenging doctrine, because it also employs the method of advocacy. The authors conclude that villages "do change in response to changes in the relative scarcity of resources, but the change takes a very long time, often generations." Rival purveyors of doctrine have little appreciation for that fact, as their interventions sometimes grossly demonstrate. I think the fundamental development assistance policy prescription which flows from this book is to proceed with caution and to check your assumptions. Village polarization increases social instability. Rapid development of markets may enhance this instability. But, greater mobility of resources and the consequent specialization of production and exchange of goods which accompanies market development is everywhere a fundamental characteristic of economic growth.

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The Agrarian Question and Reformism in Latin America

Alain de Janvry. Baltimore: The Johns Hopkins University Press, 1981, 311 pp., \$27.50 (cloth), \$8.95 (paper).

Reviewed by O. P. Blaich*

Economic development over the last three decades has been undertaken largely as a collection of disconnected projects. Therefore, the need for an integrative theory that deals with the fundamental causes of the distribution of income, wealth, and rights is one of the two major premises on which this book is based. The other is that the integrative concepts were extremely general and abstract. This situation left the practitioners of economic development in a vacuum, and they could only approach a global framework in a rather remote fashion.

The author is a professor of agriculture and resource economics at the University of California. His book draws on an extensive bibliography and is supported by his own experience and observation from extended stays in Argentina, Chile, Colombia, and Peru, as well as frequent visits to Ecuador, Mexico, Brazil, and the Dominican Republic. He examines and builds numerous theories and postulates of Marx, Ricardo, Keynes, Lenin, Kautzky, and others, and he draws on the thoughts of contemporary analysts including Prebisch, Schultz, Ruttan, and Schuh. His analysis is developed from a broad base of observation and theory, which leads him to conclude that any proposal to eliminate poverty and improve the distribution of income within and among nations needs to originate in a positive analysis of the multiple sources of unequal development. He fails to find satisfaction in neoclassical economics and chooses the perspective of political economy to integrate social, political, and economic thought—an essential element in understanding the multiple forces that underlie the demand for agrarian reform. De Janvry uses this comprehensive framework to analyze the underlying cause of change in the interdependence between large landowners and peasants, mechanization and the shift to capitalism, land reform in its various aspects, urbanization and labor availability, economic growth and the unequal distribution of income and wealth, the persistence of poverty, and other changes in the structure and performance of agriculture in Latin America.

Based on this integrated theory, de Janvry distinguishes and examines in detail the changes in agriculture during two recent periods of transformation in Latin America—the fifties, which were marked by import-substitution in industry at a time when strong vestiges of noncapitalistic or feudalistic social relations in agriculture resisted agrarian reform, and the sixties and seventies, when capitalistic modes of production were introduced to

agriculture, thereby redefining the role of the peasantry in bringing about shifts in political power and changes in social interdependence.

Within this integrative framework, de Janvry describes and analyzes these areas:

- The forces underlying the agrarian crises in several Latin American countries;
- The stagnation of food production prior to the sixties and the uneven development of capitalist and peasant agriculture in the ensuing period;
- The instability of the position of the peasant class and the dynamics of its transformation;
- The conflicts during the transition to capitalism between rents and profits, between cheap food and foreign exchange, and between peasant labor and mechanization;
- The varied nature of land reform and the transition from precapitalist to capitalist agriculture; and
- The political and economic purposes underlying some of the rural development programs and their often contradictory nature.

De Janvry concludes with a critical examination of several current policies of world agencies which purport to stimulate equitable agrarian growth. These policies include:

- The International Labor Organization proposal for labor-intensive development.
- The World Bank and the U.S. Agency for International Development's promotion of integrated rural development with emphasis on small farmers.
- The Overseas Development Council's suggestion that calls for national governments to provide an improved standard of welfare for the poor.

The author finds that each of these proposals has shortcomings as an agrarian reform strategy because it fails to deal with the fundamental causes of poverty and maldistribution of income. He subscribes to a strategy of "basic needs," but not one that would act as an income constraint on growth. Instead, he emphasizes one that would achieve those political, social, and economic structural changes needed not only to reconcile growth and distribution but also make the satisfaction of basic needs the essential purpose of economic growth. As economic growth occurs and personal income increases, the basic needs would be redefined beyond mere subsistence and would be met with rising productivity of labor.

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Although one may agree or disagree with de Janvry's final prescription for economic development, the book is well written and succeeds in developing a framework for an

analysis of agrarian issues. Whether or not it persuades the practitioners will depend on their agreement with the author's fundamental values and views.

On the Misuse of Theil's Inequality Coefficient

By Raymond M. Leuthold*

In a recent article, "The Food and Agricultural Policy Simulator," by Salathe, Price, and Gadson in this journal (4),¹ the authors use a Theil inequality coefficient (p. 7, equation (12)) to validate the performance of their model. They used actual values of Y_t and Y_{t-1} when computing this statistic. The statistic was never intended to be used in this way because it is sensitive to additive transformations. That is, if one moves the decimal point to the right for each variable, one can generate lower coefficients. Theil (5) pointed out that instead of actual values, only changes in the variables should be used. A thorough discussion of the appropriateness of Theil coefficients is contained in Leuthold (2).

Salathe, Price, and Gadson cite two references: Pindyck and Rubinfeld (3) and Kost (1). The former never mentions this problem, but the latter goes so far as to prove the additive feature. A U in actual values will always be less than a U when data are measured in changes. Kost went on to compute U, U_1 , and U_2 , and as expected, $U < U_1 < U_2$ in all but three cases. No explanation is offered for these exceptions.

Salathe, Price, and Gadson do correctly point out their statistic MARE is independent of the units of measure of Y. They go on to imply they want another statistic bounded from above by 1, but the question is: Why impose this restriction? One suspects the prime reason is to generate low U coefficients. (Again, see (3) on this point.) Their results seem to reflect the unit problem. With the exception of oats, note that the U coefficients for prices of grains (wheat, corn, barley, sorghum, soybeans, and cotton) are all higher than those for prices of meat products (barrows and gilts, steers, cows, broilers, turkeys, eggs, and milk). I suspect this occurs because all the products in the latter group are priced in units that are roughly 10

times the units of price for the former group. The U statistic would generally predict those relationships. (Milk seems to be another odd case here.) Also, the authors point out that U statistics are higher for prices than for corresponding production variables. Could that again be simply a reflection of units and not relative performance? (Perplexing here is that MARE had a similar tendency, and it is not unit-sensitive).

Salathe, Price, and Gadson have defined and used a U coefficient in a manner that Theil never intended. Theil defined it differently from the very beginning in terms of the type of data inputs. Kost used this coefficient in 1980, but he covered himself by also computing U_1 and U_2 . The U's reported by Salathe, Price, and Gadson are probably much too low, reflecting greater accuracy than is really the case. Were they to publish an addendum to their paper where they also show U_1 and U_2 , I would expect some changes in relative results among commodities, and I would expect higher coefficients reflecting the true performance errors.

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¹Italicized numbers in parentheses refer to items in the References at the end of this note.

Comment on the Misuse of Theil's Inequality Coefficient

By William E. Kost*

The unexplained inconsistency in U , U_1 , and U_2 mentioned by Leuthold in the validation results for my net trade model was due to arithmetic errors. The original calculations were done by hand because no computer programs were then available in ERS for making this calculation. When the errors were corrected, $U < U_1 < U_2$ —as Leuthold predicted.

However, there is another way an inconsistency may show up. It involves the definitions of the inequality coefficients.¹ U_1 and U_2 use lagged values of the variable. U

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¹ The Theil inequality coefficients are defined as:

$$U = \sqrt{\frac{1}{T} \sum_{t=1}^T (\hat{Y}_t - Y_t)^2} / \sqrt{\frac{1}{T} \sum_{t=1}^T \hat{Y}_t^2} + \sqrt{\frac{1}{T} \sum_{t=1}^T Y_t^2}$$

$$U_1 = \sqrt{\frac{1}{T} \sum_{t=1}^T ((\hat{Y}_t - Y_{t-1}) - (Y_t - Y_{t-1}))^2} / \sqrt{\frac{1}{T} \sum_{t=1}^T (\hat{Y}_t - Y_{t-1})^2 + \frac{1}{T} \sum_{t=1}^T (Y_t - Y_{t-1})^2}$$

$$U_2 = \sqrt{\frac{1}{T} \sum_{t=1}^T ((\hat{Y}_t - Y_{t-1}) - (Y_t - Y_{t-1}))^2} / \sqrt{\frac{1}{T} \sum_{t=1}^T (Y_t - Y_{t-1})^2}$$

where Y is the actual value, $\hat{Y}$ is the estimated value, and T is the number of periods.

does not. When using all possible data in calculating Theil inequality coefficients, the U calculation, therefore, uses one more observation than U_1 and U_2 . Depending on the magnitude of the error in this first observation, the $U \leq U_1 \leq U_2$ relationship may not hold.

The emphasis Leuthold places on the cross comparison of the three coefficients appears too strong. Although similar, these three coefficients measure different things. Their value is in pointing to ways to improve a model. Does the revised model generate a better U (or U_1 , or U_2)? The coefficients give users some idea of how good a model is. No one measure of goodness of fit is satisfactory. All measures normally used provide some useful information. The conclusion a user (or model builder) draws from the information is subjective. Though it has an additive bias, the U statistic is still a useful measure, particularly when one is comparing two versions of the same model. In that case, no scale problem exists. To use it, of course, one must know how to interpret it. But, the same is true for all other measures of goodness of fit.

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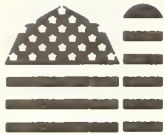
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